Information Frictions, Reputation, And Sovereign Spreads ⋆

JUAN M. MORELLI
Federal Reserve Board

MATÍAS MORETTI
World Bank

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Abstract. We formulate a reputational model in which the type of government is time varying and private information. Agents adjust their beliefs about the government's type (i.e., reputation) using noisy signals about its policies. We consider a debt repayment setting in which reputation influences the market's perceived probability of default, which affects sovereign spreads. We focus on the 2007-2012 Argentine episode of inflation misreport to quantify how markets price reputation. We find that the misreports significantly increased Argentina's sovereign spreads. We use those estimates to discipline our model and show that reputation can have long-lasting effects on a government's borrowing costs.

Keywords: Sovereign Default, Reputation, International Lending.

JEL Codes: F34, F41, G14, G15, L14

⋆ Morelli (juan.m.morellileizagoyen@frb.gov): Federal Reserve Board. Moretti (mmoretti@worldbank.org): The World Bank. We thank Mark Aguiar, Michele Cavallo, Hal Cole, Alessandro Dovis, Felicia Ionescu, Nobu Kiyotaki, Ricardo Lagos, Hanno Lustig, Pablo Ottonello, Diego Perez, Thomas Philippon, Venky Venkateswaran, and Adrien Verdelhan for useful comments. This paper was previously circulated as "Information Frictions, Partial Defaults, and Sovereign Spreads." Disclaimer: The views expressed here are our own and should not be interpreted as reflecting the views of the World Bank Group, the Board of Governors of the Federal Reserve System, or of anyone else associated with the Federal Reserve System.

1. Introduction

Policymakers usually perceive a country's reputation as an important type of gained capital to be kept over time. From a fiscal or monetary perspective, for instance, a government's history of achieving inflation or fiscal targets may affect how agents form their expectations, and thus shape the effectiveness of new policies being implemented. From a debt repayment perspective, honoring past debt obligations may affect a government's current borrowing costs and its access to different sources of credit. A crucial aspect is then quantifying how a government's reputation is affected by the policies it chooses, and how policies are, in turn, shaped by the government's reputation.

To answer these questions, we focus on a particular setting for which reputation may be a first-order concern: debt repayment. We develop a reputational model of sovereign default and provide new empirical evidence on the link between a government's reputation and its borrowing costs. The model contains multiple alternating government types, which differ in their willingness to default on their debt. Lenders do not observe the government type but use the information transmitted in the government's policies to infer it. In this context, reputation can be understood as the market belief about a government's willingness to repay given a set of macroeconomic fundamentals. Governments care about their reputation because it affects their cost of funding. The model provides a link between a government's reputation, its borrowing costs, and the policies it implements. The strength of this link depends on the way in which agents learn about the type of government from the information provided by its policies.

Guided by the model, we then go to the data and analyze a unique experiment that allows us to study the effect of a government's reputation on its borrowing costs. In particular, we focus on the Argentine 2007-2012 episode of inflation-report tampering as a case study. During 2007-2012, the Argentine government significantly underreported its inflation rate, which implied a de facto partial default on its stock of inflation-indexed bonds (IIBs). We show that the market priced the misreport, as reflected by a significant increase in the spreads of (dollar-denominated) nominal bonds. Given that coupon payments of nominal bonds were not directly affected by the misreport of inflation, we argue that the documented effects can be attributed to changes in the government's reputation.

We discipline our reputational model based on these empirical estimates. We then use the calibrated model to back out our model-implied measure of reputation and study the role of fundamentals behind the link between reputation and sovereign spreads. We show that (i) reputation matters more during "bad" states of the economy because spreads are more sensitive to reputation in those states, and (ii) changes in reputation can have long-lasting effects on borrowing costs. Finally, we bring the model to the data and show that Argentina's loss of reputation can explain 30%-50% of the increase in its sovereign spreads during the Global Financial Crisis (GFC).

Our model is in the spirit of Kreps and Wilson (1980) and Milgrom and Roberts (1982) with uncertainty about the type of government. We consider an infinite-horizon model that features incomplete markets, limited commitment, alternating government types, and noisy signals. We assume a risk-averse government that faces a stochastic endowment and issues debt in international markets. The government lacks commitment and can default on its debt. There are two types of government: a commitment type (C) and a strategic type (S). Types are time varying following a Markov process. We assume that the types differ in their incentives to default. In particular, the S-type has weakly larger incentives to default. Lenders do not observe the government type, they have a prior about the government's being of the C-type (i.e., reputation), and use the information transmitted by the government's policies to update this prior. Under this setup, changes in lenders' prior about the type of government affect their perceived default probability, and therefore the government's borrowing costs.

In addition to debt and default policies, the government can choose from a policy $\tilde{\pi}$ that provides a benefit in terms of current consumption. We assume that the S-type can choose any value for $\tilde{\pi}$ but the C-type commits to $\tilde{\pi} = 0$. Although there are no direct costs associated with this policy, by setting $\tilde{\pi} \neq 0$ the S-type may signal its type, which affects its borrowing costs. Lenders do not perfectly observe the policy $\tilde{\pi}$ but receive a noisy signal about it, which implies that they only learn from it gradually. We interpret this policy as any action that can potentially provide information about the type of government. For the Argentine case, for instance, $\tilde{\pi}$ can be interpreted as the inflation misreport policy, since it may be informative about the government's willingness to default.

We then use the Argentine 2007-2012 episode of inflation misreport to analyze the effects of a government's reputation on its borrowing costs. During these years, the official Consumer Price Index (CPI) was intentionally underreported by the national government (see Cavallo, 2013 and Cavallo et al., 2016 for a detailed discussion). We focus on this episode for the following reasons. First, the misreports were large and significantly affected coupon payments of IIBs. During this episode, the amount outstanding of Argentina's IIBs accounted for almost a quarter of its stock of debt, so that the underreport of inflation had a great impact on the government's stock of debt. Second, the misreports occurred frequently, which allows us to work with a large number of observations. Third, Argentina was not excluded from international debt markets as a consequence of this policy. We can then use secondary markets data to quantify the contemporaneous effect of the misreports on Argentina's spreads. Lastly, the misreport only affected coupon payments of IIBs. By studying the effects of this policy on other types of bonds (e.g., nominal bonds), we can then isolate the reputational effects of the misreport.

There are two main challenges in assessing the causal effect of inflation tampering on Argentina's spreads. The first is measurement, given that lenders cannot perfectly observe the "true" inflation rate and hence the magnitude of the misreport.[1] Moreover, based on our reputational model, only unexpected changes in the misreport should have an effect on prices. If the market was expecting the misreport, that effect should already be priced. To address this concern, we consider changes in the break-even (BE) inflation rate as a proxy for the unexpected misreport.[2] Embedded in the BE inflation rate is the market's expectation about the inflation announced by the government, since these announcements directly affect the returns of IIBs. Changes in the BE rate around days on which the government reported the inflation rate can therefore be used to infer the market's surprise.

The second challenge is reverse causality, since inflation tampering may be the government's response to a rise in spreads. If that is the case, a simple OLS regression would yield biased point estimates. To address this concern, we adopt a heteroskedasticity-based identification strategy (Rigobon and Sack, 2004) and exploit changes in the volatility of the BE inflation rate around days on which the government reported the inflation rate. The main identifying assumption is that the volatility of shocks to the BE inflation rate is significantly higher around these announcements, but the variance of shocks to sovereign spreads (and other common shocks) remains the same.

We show that the sequence of misreports significantly increased the spreads of dollar-denominated bonds issued by the Argentine government. In particular, we find that a 1–sd decrease in the BE inflation rate leads to a rise in spreads that accounts for more than two thirds of their daily dispersion. Interpreted through the lens of our reputational model, given that coupon payments of dollar-denominated bonds were not directly affected by the misreports,

[1] Cavallo et al. (2016) show that the lack of reliable official data led to the creation of several unofficial inflation indicators. Agents can then use these alternative indices to get a noisy signal about the magnitude of the misreport.

[2] The BE inflation rate is the level of inflation that renders an investor indifferent between holding nominal bonds or IIBs.

these results suggest that a government's reputation can play an important role in the pricing of sovereign bonds.

For the quantitative analysis, we discipline our reputational model based on these empirical estimates. In particular, we use our estimated (semi) elasticity to pin down how agents learn about the type of government through the information provided by the policies it chooses. Since our empirical estimates rely on high-frequency market reactions, we provide a simple extension of our baseline model to include secondary markets. In this way, we can capture the intraperiod effect of $\tilde{\pi}$ on sovereign spreads, as we do in the data. We then use the calibrated model to compute a measure of reputation and to assess the role of fundamentals behind the link between reputation and sovereign spreads. We show that reputation matters more during bad states of the economy. This is because in bad economic times, spreads are significantly more sensitive to lenders' beliefs, which resembles the result in Cole and Kehoe (2000). Finally, we bring the model to the data and show that changes in reputation can have long-lasting effects on borrowing costs. In particular, we find that Argentina's loss of reputation can explain up to 30%-50% of the increase in its sovereign spreads during the GFC.

Literature Review

Our paper relates to a large literature on how the presence of asymmetric information about a government's type affects its policies and different macroeconomic outcomes. Backus and Driffill (1985); Barro (1986); Persson and Tabellini (1997); Phelan (2006); and Dovis and Kirpalani (2020) examine the role of a government's reputation in the design of fiscal, monetary, and regulatory policies. In particular, our paper contributes to a growing body of work that studies reputation dynamics when players' actions are not perfectly observable (Cripps, Mailath, and Samuelson, 2004; Ekmekci, 2011; Faingold and Sannikov, 2011; Board and Meyer-Ter-Vehn, 2013; Faingold, 2020; Bohren, 2021). A close study in this regard is Dovis and Kirpalani (2021), who analyze the optimal transparency of governments' rules in a context in which the type of government is private information. We contribute to this literature by providing a framework that links a quantitative analysis of the role of a government's reputation with a relevant empirical counterpart.

Our paper contributes to the literature on sovereign defaults and governments' reputation. Close studies in this area are Cole, Dow, and English (1995); Alfaro and Kanczuk (2005); D'Erasmo (2011); Amador and Phelan (2021); and Fourakis (2021). As in our study, these papers analyze a sovereign debt model with limited commitment à la Eaton and Gersovitz

(1981), in which the type of government is time varying and private information.[3] We contribute to this literature by providing new empirical evidence on the link between a government's reputation and its borrowing costs. We then use those estimates to calibrate our reputational model. In particular, we discipline the way in which agents learn about the government type from the information transmitted by its policies. Using the calibrated model, we quantify the incidence of a government's reputation on its borrowing costs.

Our paper is related to a large empirical literature that estimates the effects of a government's history of outright defaults on its borrowing costs (see, for example, Özler, 1993; English, 1996; Reinhart et al., 2003; Borensztein and Panizza, 2009; Cruces and Trebesch, 2013; Benczur and Ilut, 2016; and Catao and Mano, 2017). A shortcoming of these papers is that outright defaults are infrequently observed in the data and typically lead to exclusion from debt markets, which makes it hard to identify the effects of reputation.[4] Moreover, given that a sovereign default usually takes a long time to resolve, a government's default history may not be a good predictor of its current reputation. We address these shortcomings by focusing on an episode of recurrent partial defaults (i.e., the misreports) and by providing a high-frequency identification strategy using financial markets data.

Our high-frequency identification strategy is closely related to that of Bernanke and Kuttner (2005); Rigobon and Sack (2004); and, particularly, Hébert and Schreger (2017). Our work contributes in this dimension by estimating the short-run effect of Argentina's inflation misreport on its sovereign spreads. We argue that the documented effects are mainly due to changes in the government's reputation, and provide a quantitative model to formalize the mechanism.

Lastly, our paper is related to a quantitative literature on sovereign partial defaults. Arellano et al. (2022) provide a model in which a government can partially default on its debt obligations directly. Aguiar et al. (2013); Phan (2017a); Ottonello and Perez (2019); Du and Schreger (2022); and Engel and Park (2022) formulate models in which a government can partially default on its stock of nominal bonds by increasing the inflation rate. All of these studies assume either an exogenous output loss or exclusion from debt markets as a punishment for a

[3]Another related paper is Cole and Kehoe (1998), in which the government type is private information but fixed. In turn, other studies, such as Sandleris (2008); Phan (2017b); and Dovis (2019), analyze models in which the type of government is public information, but in which the government uses debt and default policies as a signaling device about the economy's fundamentals.

[4]These studies do not disentangle whether the rise in sovereign spreads after a sovereign default can be attributable to a punishment or reputational effect. The exception is Benczur and Ilut (2016), who pose a structural-form asset-pricing regression to disentangle the role of reputation.

partial default.[5] We contribute to this literature by providing a microfoundation for the costs of partial defaults, based on a government's reputation in international debt markets.

The rest of the paper is structured as follows. Section 2 presents the reputational sovereign default model with noisy signals. Section 3 describes the empirical analysis, based on Argentina's inflation-tampering episode. Section 4 presents the quantitative analysis, and Section 5 concludes.

2. A Reputational Model of Sovereign Default

2.1. *Model Description*

We consider a small open economy with incomplete markets that receives a stochastic endowment y, which follows a continuous Markov process with a transition function $f(y' \mid y)$. An infinite-lived risk-averse government issues debt in international markets to smooth its consumption.

There are two types of government: a commitment type (C-type) and a strategic opportunistic type (S-type). We assume that the government type exogenously changes over time, based on a stochastic Markov process denoted by T.[6] Government types differ in their incentives to default. In the spirit of Kreps and Wilson (1980) and Milgrom and Roberts (1982), the type is not publicly observable. Lenders have a prior ζ about the government's being of the C-type, which they update based on the information transmitted by the government's policies.

The government issues long-term non-contingent bonds, b. We assume debt contracts that mature probabilistically, as in Chatterjee and Eyigungor (2012). Each unit of b matures in the next period with probability λ. If the bond does not mature and the government does not default, it pays a coupon z. The government lacks commitment and can default on its debt obligations. Let $d = \{0, 1\}$ be the outright default policy on b, where $d = 1$ denotes a default. As is standard in the literature, an outright default leads to a temporary exclusion from debt markets and an exogenous output loss, $\phi_j(y)$. We assume that $\phi_C(y) \geq \phi_S(y)$ for all

[5] The exception is Du and Schreger (2022). In this case, the cost is endogenous and depends on the foreign currency mismatch on corporate balance sheets.

[6] A well-known result of Cripps, Mailath, and Samuelson (2004) for the context of repeated games is that in a model with *fixed* types and imperfect monitoring, reputation is a short-run phenomenon. Any model of long-run reputation should thus include some mechanism by which the uncertainty about types is continually replenished. See Ekmekci (2011); Board and Meyer-Ter-Vehn (2013); or Bohren (2021) for different ways in which the uncertainty can be replenished.

y, meaning that the S-type has (weakly) larger incentives to default.[7] Changes in lenders' prior about the type of government thus affect their perceived probability of default, and therefore the government's borrowing costs.

In addition, the government can decide on another policy $\tilde{\pi} \leq 0$, which provides a benefit of $\Omega(\tilde{\pi})$ in terms of additional consumption c. We assume that $\Omega(\tilde{\pi})$ is increasing in $\mid \tilde{\pi} \mid$. The S-type can choose any value $\tilde{\pi} \leq 0$, but the C-type commits to $\tilde{\pi} = 0$. The policy does not lead to a direct cost (such as exclusion from markets or output losses). However, it provides information about the type of government. We assume that $\tilde{\pi}$ is not perfectly observable by lenders. Instead, they receive a noisy message m, whose realization depends on the government's choice for $\tilde{\pi}$.[8] For tractability, we assume that the noisy message takes two values, $m = \{L, NL\}$, where L (lie) signals $\tilde{\pi} \neq 0$.[9] The probability of receiving message L is given by

$$Prob\,(m = L \mid \tilde{\pi}) = \Gamma\left(\tilde{\pi}; \sigma, \alpha\right), \tag{1}$$

where the parameter $\sigma \geq 0$ captures the noise behind the underlying message and $\alpha \leq 0$ is a learning parameter that governs how agents learn from this policy. We assume that $\Gamma\left(\cdot\right)$ is increasing in the magnitude of $\tilde{\pi}$ (i.e. $\Gamma'_{\tilde{\pi}}\left(\cdot\right) < 0$) and (weakly) increasing in α. This implies that (for a given noise σ) agents can more easily detect $\tilde{\pi} \neq 0$ as α increases.[10]

Figure 1 describes our timing assumption. Let $\mathbf{S} = (y, b, \zeta)$ be the state at the beginning of the period. Each period is divided into 3 stages. In stage 0, the government chooses to default or not ($d = \{0, 1\}$) on b. Lenders observe this action and update their beliefs accordingly ($\tilde{\zeta}$). If the government defaults, it faces an output cost $\phi_j(y)$ and is temporarily excluded from international debt markets. We assume that it regains access to debt markets with probability θ in the next period. There is no recovery value and the stock of debt is $b = 0$ after exiting a default.

7In Alfaro and Kanczuk (2005) and D'Erasmo (2011), the types differ in their discount factor. Our specification is similar to the one in Barret (2016) or Egorov and Fabinger (2016), and it can be interpreted as differences in the disutility over an outright default (as in Cole and Kehoe, 1998).

8In this regard, our study is similar to that of Holmström (1999) and Mailath and Samuelson (2001) because it features both noisy signals and alternating types.

9We take this notation motivated by the Argentine case, in which the government was either lying or not about the inflation rate.

10A realization of $m = L$ does not necessarily reveal that the government is of the S-type, since we assume that $\Gamma\left(0; \sigma, \alpha\right) \geq 0$. Under $(\sigma, \alpha) = (0, 0)$, the message $m = L$ is perfectly informative.

FIGURE 1. Timing of Events

	If default	If no default	
Stage 0	Stage 1	Stage 1	Stage 2
- Initial $\mathbf{S} = (y, b, \zeta)$ - Default choice $d = \{0, 1\}$ - First update of beliefs $\tilde{\zeta}(d, \zeta)$	- Temporary exclusion from debt markets - Output cost $\phi_j(y)$	- Choice of b' and $\tilde{\pi}$ - Message m is realized - Second update of beliefs $\hat{\zeta}(m, \tilde{\zeta})$	- Debt issuance b'

If the government does not default, then choices for $\tilde{\pi}$ and b' are made at stage 1. We assume that both the C- and S- types follow the same debt policy, $b^{\star\prime}\left(y, b, \tilde{\zeta}\right)$. We interpret this policy as a fiscal rule that is not under the control of the j-type. Instead of imposing an arbitrary fiscal rule, we assume that bond policies are optimally chosen by another agent of the economy (say, the Congress), whose information set is the same as that of the lenders. Under this assumption, bond policies are uninformative about the type of government. An advantage of this specification is that it allows us to compare our model with others in the sovereign debt literature. For instance, if we assume that the type of government is fixed and publicly known, then the bond policy $b^{\star\prime}\left(y, b, \tilde{\zeta}\right)$ would be exactly the same as that in Chatterjee and Eyigungor (2012).

Given the choice of $\tilde{\pi}$, message m is realized and lenders once again update their beliefs $(\hat{\zeta})$ at the end of stage 1. At stage 2, the primary market for bonds opens and the government issues b' (chosen at stage 1), taking the bond price schedule $q(\cdot)$ as given.[11] Under this setup, the resource constraint of the economy is given by

$$\begin{aligned} c(d=0, \tilde{\pi}, b^{\star\prime}) &= y - b\left[(1-\lambda)\, z + \lambda\right] + q\left(\cdot\right)\left[b^{\star\prime} - (1-\lambda)b\right] + \Omega(\tilde{\pi}) \\ c(d=1) &= y - \phi_j(y). \end{aligned} \tag{2}$$

2.2. *Noisy Signals, Update of Beliefs, and Bond Prices*

As shown in the timeline of Figure 1, beliefs about the government type are updated twice within a period: After the outright default decision d and after the message m is realized. Let $d_j^\star \equiv d_j^\star\left(y, b, \zeta\right)$ be the lenders' conjecture about the j-type government's default decision.

[11]Under our timing assumption, b' is chosen before message m is realized. This allows us to isolate the effect of the message on the government's reputation and spreads (see Section 4.1).

Based on Bayes' rule, the first updating of beliefs is given by[12]

$$\tilde{\zeta}\left(d, \zeta; d_S^\star, d_C^\star\right) = \frac{Prob(d \mid d_C^\star) \times \zeta}{Prob(d \mid d_C^\star) \times \zeta + Prob(d \mid d_S^\star) \times (1-\zeta)}. \tag{3}$$

If the government did not default, the second updating of beliefs happens after the j-type chooses $\tilde{\pi}$ and lenders observe the message m. Let $\tilde{\Pi}_j^\star \equiv \tilde{\Pi}_j^\star(y, b, \tilde{\zeta})$ be the lenders' conjecture about the j-type's $\tilde{\pi}$ policy. For a given realization of m, the updated beliefs are given by[13]

$$\hat{\zeta}\left(m, \tilde{\zeta}; \tilde{\Pi}_S^\star, \tilde{\Pi}_C^\star\right) = \frac{Prob(m \mid \tilde{\Pi}_C^\star) \times \tilde{\zeta}}{Prob(m \mid \tilde{\Pi}_C^\star) \times \tilde{\zeta} + Prob(m \mid \tilde{\Pi}_S^\star) \times \left(1-\tilde{\zeta}\right)}. \tag{4}$$

Taking into account the Markov transition across the two government types (T), the end-of-period posterior is given by

$$\zeta'\left(\hat{\zeta}\right) = T_{CC} \times \hat{\zeta} + T_{SC} \times \left(1-\hat{\zeta}\right). \tag{5}$$

Equations (1), (4), and (5) imply that, through its effects on m, changes in $\tilde{\pi}$ affect the government's reputation ζ'. Panel (A) of Figure 2 provides a graphical illustration. Once the message m has been realized (and given the lenders' conjectures), ζ' is independent of the current choice of $\tilde{\pi}$ (horizontal solid lines). Ex ante, however, a larger $|\tilde{\pi}|$ increases the probability that message $m = L$ is realized, which affects the expected ζ' (dashed lines). The effect depends on how agents can learn from the policy $\tilde{\pi}$. For instance, a larger α increases the probability of message $m = L$ being realized for any $\tilde{\pi} \neq 0$, which affects the sensitivity of a government's reputation to $\tilde{\pi}$.[14]

We assume that lenders are risk neutral. The price of a bond is thus given by the expected value of repayment, discounted by the risk-free rate r. Let $VR_j\left(y', b', \zeta'\right)$ be the next-period value of repayment if the government is of the j-type. The bond-pricing kernel is given by

$$q\left(y, b', \zeta'\right) = \frac{1}{1+r}\int_y \left\{\zeta' VR_C\left(y', b', \zeta'\right) + \left(1-\zeta'\right) VR_S\left(y', b', \zeta'\right)\right\} dF\left(y' \mid y\right) \tag{6}$$

with

$$VR_j\left(y', b', \zeta'\right) \equiv \left(1 - d_j^{\star\prime}\right) \times \left[\sum_{M=\{L,NL\}} Prob(m' = M \mid \tilde{\Pi}_j^{\star\prime})\left(\lambda + (1-\lambda)\left[z + q'_M\right]\right)\right], \tag{7}$$

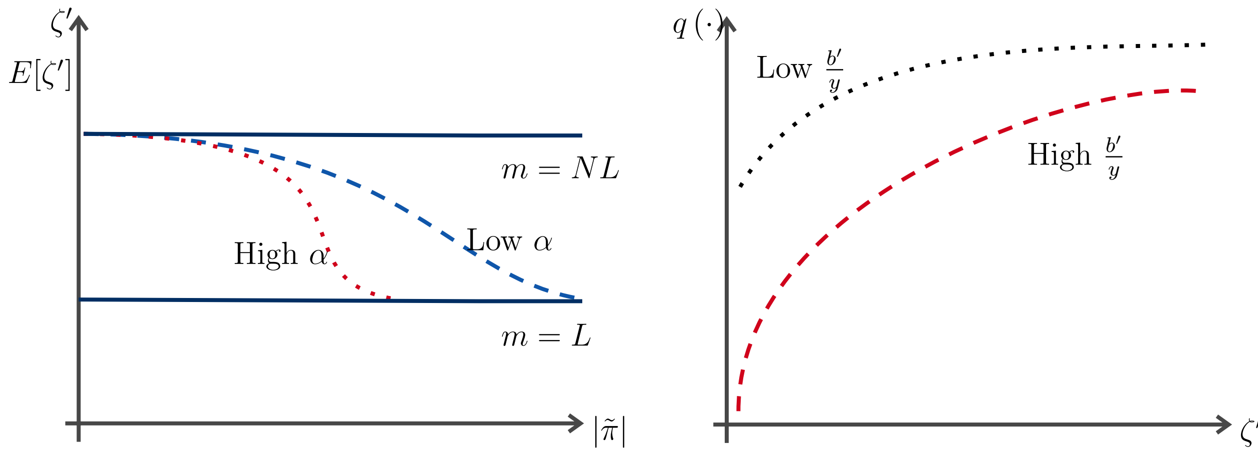
[12] For off-equilibrium paths, we simply assume that $\tilde{\zeta}\left(d, \zeta; d_S^\star, d_C^\star\right) = 0$.

[13] Regardless of the choice of $\tilde{\pi}$, we assume that both messages have positive probability, so Bayes' rule always applies and there are no off-path information sets.

[14] From Equations (1), (4), and (5), changes in α also affect ζ' and thus the horizontal lines of Figure 2. In Appendix C.2, we provide a detailed analysis of the channels through which α affects the expected posterior.

FIGURE 2. Noisy Signals, Reputation, and Bond Prices

(A) $\tilde{\pi}$ and Reputation (B) Reputation and Prices



Notes: Panel (A) shows the realized and expected posteriors as a function of $|\tilde{\pi}|$ for a given $\tilde{\zeta}$. The top and bottom horizontal lines depict ζ' when $m = NL$ and $m = L$ are realized, respectively. The dashed lines show the expected posterior, $\mathbb{E}[\zeta']$, for two values of α. Panel (B) shows the pricing kernel as a function of ζ' for different values of (y, b').

where $d_j^{\star\prime}$ and $\tilde{\Pi}_j^{\star\prime}$ refer to the conjectured next-period policies for type j. The term q_M' refers to the next-period price for one unit of debt. This price is also a function of lenders' conjectures and is contingent on the realization of the next-period message (see Appendix A.2).

Under the assumption that $\phi_C(y) \geq \phi_S(y)$ for all y, the S-type has weakly larger incentives to default on b. This implies that $VR_C(y', b', \zeta') \geq VR_S(y', b', \zeta')$, and thus bond prices are weakly increasing in ζ'. The effects are state-contingent, since they depend on the economy's fundamentals y and b. Panel (B) of Figure 2 illustrates this point. The figure shows the pricing kernel $q(y, b', \zeta')$ as a function of ζ', for two ratios of b'/y. When b'/y is small (dotted line), the default probability for both the C- and the S-type defaults is low. This implies a small difference between $VR_C(\cdot)$ and $VR_S(\cdot)$, and thus changes in reputation have a small effect on the pricing kernel $q(\cdot)$.

2.3. *Government's Recursive Problem*

We briefly describe the government's recursive problem, focusing on the S-type optimal choice of $\tilde{\pi}$. We leave a detailed description of the problem to Appendix A.1.

If the government is not in default, the beginning-of-period value function, $W_j(y, b, \zeta)$, depends on the optimal default decision at stage 0. For the j-type, it is given by

$$W_j(y, b, \zeta) = \max_{d \in \{0,1\}} \left\{ W_j^R\left(y, b, \tilde{\zeta}\right), W_j^D\left(y, \tilde{\zeta}\right) \right\}, \tag{8}$$

where $W_j^R(\cdot)$ denotes the value function in case of repayment, $W_j^D(\cdot)$ is the value function in case of default, and $\tilde{\zeta}$ is given by Equation (3). The value of default depends on the output cost $\phi_j(y)$ and the probability of exiting the default status θ. In this section, we describe the $W_j^R(\cdot)$ function and in Appendix A.1 we describe $W_j^D(\cdot)$.

At stage 1, taking as given the bond policy rule $b^{\star\prime} \equiv b^{\star\prime}\left(y, b, \tilde{\zeta}\right)$, the S-type solves for the optimal $\tilde{\pi}$ policy. In particular, it chooses $\tilde{\pi} \in [\underline{\pi}, 0]$ to maximize the weighted average of the value function in stage 2, $V_S(\cdot)$, where the weights are given by the probability that message m is realized, given the choice of $\tilde{\pi}$. The problem is as follows:

$$W_S^R\left(y, b, \tilde{\zeta}\right) = \max_{\tilde{\pi}} \sum_{M=\{L,NL\}} Prob(m = M|\tilde{\pi}) \times V_S\left(\tilde{\pi}, y, b, \hat{\zeta}(M)\right) \tag{9}$$
$$s.t. \quad \tilde{\pi} \in [\underline{\pi}, 0],$$

where $\hat{\zeta}(m)$ is the posterior defined in Equation (4) and $V_S(\cdot)$ is given by

$$V_S\left(\tilde{\pi}, y, b, \hat{\zeta}(m)\right) = u(c) + \beta \int_y \Big\{ T_{SS} W_S(y', b^{\star\prime}, \zeta') + T_{SC} W_C(y', b^{\star\prime}, \zeta') \Big\} dF(y' \mid y) \tag{10}$$
$$\text{s.t. } c = y - b\left[(1-\lambda) z + \lambda\right] + q(y, b^{\star\prime}, \zeta')\left[b^{\star\prime} - (1-\lambda)b\right] + \Omega(\tilde{\pi}),$$

where β is the government's discount factor and ζ' is given by Equation (5). The S-type, thus, faces a stochastic trade-off when choosing the optimal $\tilde{\pi}$. Conditional on the realization of message m, since $\Omega(\tilde{\pi})$ is increasing in the magnitude of $\tilde{\pi}$, the value function $V_S(\cdot)$ is increasing in $\mid \tilde{\pi} \mid$. A larger $\mid \tilde{\pi} \mid$, however, increases the probability that message $m = L$ is realized, which decreases $\hat{\zeta}$ and raises borrowing costs.

2.4. *Link with the Argentine Case*

The previous model provides a mapping from a government's policies to its reputation, and from reputation to bond prices. Underlying this mapping is the way in which agents can learn from those policies; in particular, from $\tilde{\pi}$. The policy $\tilde{\pi}$ can be interpreted as any government action that signals its type. Based on our Argentine case of study, we will interpret this policy as a misreport of the inflation rate that dilutes the real value of inflation-indexed bonds (IIBs). Under this interpretation, the parameter α determines how agents learn about the type of government based on the sequence of misreports.

We assume that the government faces a constant legacy stock of IIBs, whose coupon payments are linked to the inflation announced by the government. For tractability, we assume that this debt is a perpetuity and we denote its coupons with B. The S-type can affect coupon payments

B by underreporting the inflation rate and choosing $\tilde{\pi} \in [\underline{\pi}, 0]$. Under this setup, notice that $\tilde{\pi} < 0$ implies an indirect partial default on B.

If not in default, the resource constraint of the economy can be written as

$$c(d = 0, \tilde{\pi}, b^{\star\prime}) = y - b\left[(1-\lambda)\, z + \lambda\right] + q\left(\cdot\right)\left[b^{\star\prime} - (1-\lambda)b\right] - B \times (1 + \tilde{\pi})\,.$$

In Section 3, we use the Argentine episode of inflation misreport to infer the sensitivity of $q(y, b', \zeta')$ to changes in a government's reputation ζ'. To this end, we use high-frequency market reactions to estimate the elasticity of q to changes in $\tilde{\pi}$. In Section 4, we then use those estimates to discipline our quantitative model. In particular, we use the empirical elasticity to pin down the learning parameter α, which links the misreports with changes in the government's reputation. Since our empirical estimates rely on high-frequency market reactions, we provide a simple extension of our baseline model to include secondary markets. In this way, we can capture the intraperiod effect of $\tilde{\pi}$ on q. This extension nests the baseline model and is described in Appendix A.4.

3. Empirical Analysis: The Case of Argentina

In this section, we provide evidence on the effect of a government's reputation on its borrowing costs. To this end, we use the Argentine 2007-2012 episode of inflation misreport as a case study. During this period, the official CPI was intentionally underreported by the national government. The sequence of misreports directly affected the coupon payments of IIBs and can therefore be interpreted as an indirect partial default on these bonds.

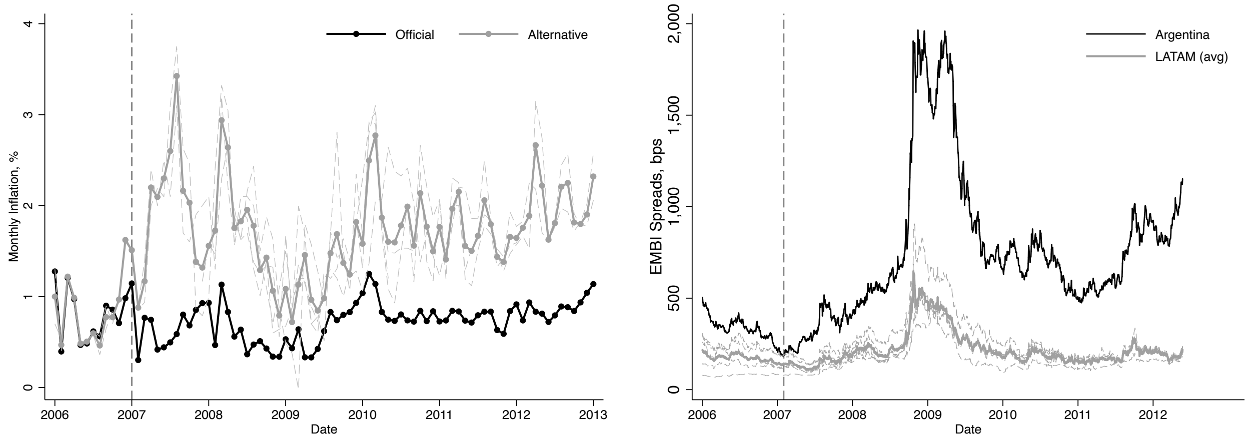
We focus on the Argentine government's systematic misreport of inflation for the following reasons. First, the misreports were large and significantly affected coupon payments of IIBs. During this period, the amount outstanding of Argentina's IIBs accounted for almost a quarter of its debt according to official data.[15] By lowering interest payments and principal, the underreport of inflation had a great impact on the government's stock of debt and implied an indirect partial default on the stock of IIBs.[16] Second, the misreports occurred frequently, allowing us to work with a relatively large number of observations. Third, Argentina was not excluded from international debt markets as a consequence of this policy. We can then use secondary markets data to quantify the contemporaneous effect of the misreports on Argentina's spreads. Lastly,

[15] See https://www.argentina.gob.ar/economia/finanzas/deudapublica/informes-trimestrales-de-la-deuda.

[16] By misreporting its inflation rate, Argentina decreased its IIB payments by nearly \$3.2 billion, which accounts for around 1% of its GDP.

Figure 3. Argentina's Misreport of Inflation and Decoupling of Spreads

(a) Argentina's measures of inflation (b) Latin America's bond spreads



Notes: The left panel shows the monthly official inflation rate announced by the Argentine government (black line) and alternative measures of inflation (gray lines). The right panel shows annualized EMBI spreads for Argentina (black line) and for other Latin American countries (gray lines). Vertical lines denote the first month in which the Argentine government underreported the inflation rate.

coupon payments of dollar-denominated bonds were not directly affected by the misreport of inflation. This allows us to isolate the reputational effect of such policy.

For most of the first half of the 2000s, Argentina's inflation rate was relatively low compared to its historical values but it peaked in 2005 at more than 10%.[17] The response of the government was to impose a series of price controls in 2006 and to pressure the staff of the National Statistics Institute (INDEC) to manipulate the national price index computed in that institution. In February 2007, the government directly intervened with the INDEC and fired its highest ranked members, including the statistician in charge of elaborating the CPI.[18]

The left panel of Figure 3 shows the announced inflation rate for the period under analysis. The reported inflation was consistently lower than other (private) measures of inflation, which we regard as noisy signals for market participants. The magnitude of the underreport—the difference between alternative measures and the official measure—was sizable and persistent.

The right panel of Figure 3 shows that in tandem with the government's systematic misreport of inflation, the Argentine spreads for dollar-denominated bonds started to decouple from those of the rest of Latin America. This is surprising for at least three reasons. First, Argentina's

[17]The average annual inflation rate for 1984-2004 was 74% and the median rate was 11.4%. In contrast, the average annual inflation rate for 2000-2004 was 7.6% and the median was 3.5%.

[18]See Cavallo et al. (2016) for a complete timeline of all events from 2006 to 2015.

fundamentals were in line with those of other Latin American countries.[19] Second, the coupons for dollar-denominated bonds were not directly affected by the misreport of inflation. Third, by underreporting the inflation rate, the Argentine government significantly decreased the real value of its stock of IIBs. In the absence of a reputational type of channel, the lower real stock of debt should decrease the spreads of nominal bonds denominated in dollars.[20] In what follows, we measure the extent to which this increase in spreads can be attributed to the inflation misreport and provide evidence in favor of a reputational channel.

3.1. *Identification Strategy*

Our main hypothesis is that the underreporting of inflation is informative for lenders regarding the government's willingness to default on its obligations, and should then affect sovereign spreads. There are, however, two main challenges to the identification of this effect: (i) measurement and (ii) reverse causality. The former arises because the government's misreport is not directly observable. The latter arises because the misreport may be a government's best response to a deterioration of the economy's fundamentals.

The first main challenge is to quantify the unexpected part of the misreport. To the extent that agents had anticipated the underreport, the government's announcement of inflation does not provide the market with additional information and sovereign spreads should not react to that announcement. In other words, only unexpected movements in the misreport provide information to agents. Our premise is that changes in the break-even inflation rate (ΔBE_t) around days on which the government announces the inflation rate can be used as a proxy for the unexpected misreport.

The break-even rate is the level of inflation that renders an investor indifferent between holding nominal bonds or IIBs. It can be computed as $BE_t = \text{Yield}_t^{\$} - \text{Yield}_t^{\text{IIB}}$, where $\text{Yield}_t^{\$}$ is the yield of a nominal bond denominated in local currency (pesos) and $\text{Yield}_t^{\text{IIB}}$ is the yield of an inflation-linked bond with similar maturity. Embedded in BE_t is the market's expectation regarding the inflation announced by the government, since these announcements directly affect the return of IIBs.[21] The main advantage of using ΔBE_t is that it is a high-frequency variable

[19] In Appendix B.3, we provide some figures to show that if anything, GDP growth in Argentina was higher than the average growth rate for the region. Argentina's stock of external debt, moreover, displayed a downward trend during this period.

[20] In canonical models of sovereign default (e.g., Arellano, 2008; Chatterjee and Eyigungor, 2012), spreads are typically increasing in the stock of government debt.

[21] This is because the coupon payments of IIBs are directly linked to the inflation reported by the government. The argument implicitly assumes a frictionless market. The BE rate may also reflect a liquidity or risk premium

that allows us to focus on narrow windows around inflation announcements. The day before the government's announcement of inflation (i.e., at time $t-1$), absent a liquidity-premium component, we should expect $BE_{t-1} \simeq \mathbb{E}_{t-1}(\hat{\pi}_t)$, where $\mathbb{E}_{t-1}(\hat{\pi}_t)$ is the market's expected announcement at time t. After the government reports $\hat{\pi}_t$, the change in the BE rate should thus be close to $\Delta BE_t \simeq \hat{\pi}_t - \mathbb{E}_{t-1}(\hat{\pi}_t)$.

Changes in the break-even inflation rate allow us to capture the difference between the government's announced inflation rate and the announcement expected by the market. However, in the Argentine case, there are two different components behind ΔBE_t: (i) the unexpected misreport and (ii) news about the "true" inflation rate.[22] In Subsection 3.4 and in Appendix B.5, we present evidence that suggests that changes in BE_t around days on which the government reported the inflation rate were mainly driven by changes in the unexpected misreport.

Apart from measurement, another concern is that agents may learn through time about the type of government. If the government's reputation ζ is already deteriorated, then an unexpected misreport should not have a significant effect on bond prices. The sequence of misreports in early 2007 may thus have different implications compared with the sequence of misreports in 2010. To overcome this concern, we split our analysis across different years. For our main specification, we focus on the period between the first misreport of inflation (January 2007) and the beginning of the Global Financial Crisis. We take the collapse of Bear Stearns on March 13, 2008 as the start of the crisis. We then study the effects for later periods and provide evidence in favor of the hypothesis that agents learned about the type of government through time.

A second challenge behind the identification is reverse causality. That is, the underreport of inflation may be the government's optimal response to a change in sovereign spreads, SP_t, due to a worsening of fundamentals. In addition, there may be (potentially unobserved) common shocks that drive, at the same time, changes in BE_t and SP_t.[23]

To address these concerns, we adopt a heteroskedasticity-based identification strategy similar to the one used in the monetary policy literature to identify the effects of monetary policy shocks

component. To the extent that this premium is constant across time, changes in the BE rate are still a good proxy for the unexpected misreport of inflation.

[22]To see this, assume that the announced $\hat{\pi}$ can be decomposed in a "true" inflation component, π, and a misreport component, $\tilde{\pi}$. Then, $\Delta BE_t \simeq (\pi_t - \mathbb{E}_{t-1}\pi_t) + (\tilde{\pi}_t - \mathbb{E}_{t-1}\tilde{\pi}_t)$, where the first term is the surprise regarding true inflation and the second term is the surprise regarding the misreport.

[23]Examples of these common factors are changes in risk aversion, flight-to-liquidity, or flight-to-safety type of events.

(Rigobon and Sack, 2004). In particular, we exploit high-frequency changes in the volatility of ΔBE_t around days on which the government announced the inflation rate.

This type of identification allows us to tackle both the reverse causality and common factors concerns. First, by focusing on changes in BE_t in narrow windows around the inflation announcement, we can ameliorate the concern that the misreport was an optimal response to an increase in SP_t. This is because the process of measuring and announcing the inflation rate takes time (even if it is not correctly measured), and it is therefore unlikely that the current (daily) change in SP_t is behind the announced inflation. Moreover, unlike an event-study analysis, the heteroskedasticity-based identification strategy does not require the complete absence of common shocks—an assumption that may be too strong in our setup. Instead, it relies on the weaker assumption that the volatility of these shocks remains constant around days on which the government announced the inflation rate.

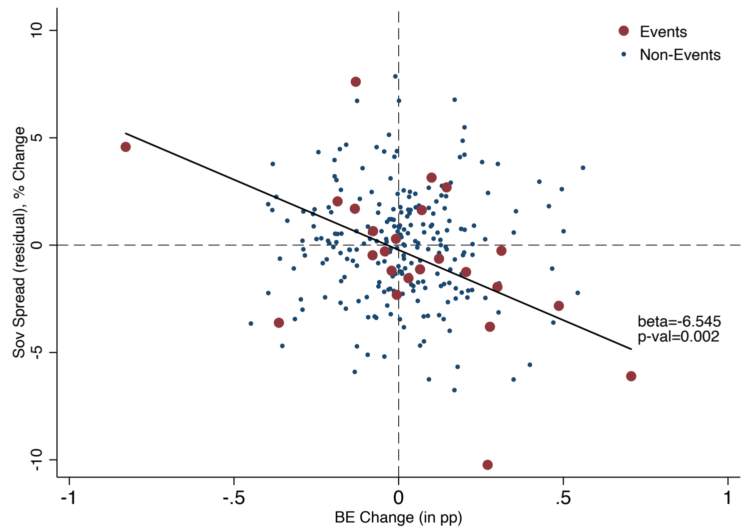
3.2. *Data and Summary of Events*

We use the J.P. Morgan EMBI spread as a measure of the Argentine government's spreads. This index captures spreads for bonds denominated in foreign currency. We use changes in the break-even inflation rate as a proxy for the unexpected misreport of inflation, as explained in Section 3.1. A problem with the Argentine case during the period of study is the lack of bonds denominated in local currency, which are needed to construct the BE inflation rate.[24] To circumvent this issue, we use dollar-denominated bonds, adjusting their yields using the expected depreciation rate of the Argentine peso implied by currency forward contracts. Appendix B.4 provides details on the construction of the BE inflation rate. Appendix B.5 analyzes the relation between changes in the BE rate and the unexpected misreport. Appendix B.6 discusses the role of the exchange rate in our constructed measure.

For our baseline analysis, we focus on the period between January 2007 and February 2008. Figure 4 shows the relation between ΔBE_t and changes in Argentina's sovereign spreads after controlling for global factors. Red dots indicate 2-day windows around days on which the Argentine government reported the inflation rate; see Appendix B.2 for the full list of days. We name these days *event days* (E). During event days, the relation between ΔBE_t and $\Delta \ln SP_t$ is negative and significant, indicating that an increase in the unexpected underreport of inflation (i.e., a lower $\Delta BE_t < 0$) is associated with a higher increase in sovereign spreads. All other

[24]There is only one bond denominated in pesos for which we have data during our sample period, and the first observation is for July 2007—i.e., 6 months after the government started misreporting the inflation rate.

FIGURE 4. Break-even Inflation Rate



Notes: The figure shows the daily change in BE_t and the daily log change in Argentina's sovereign spreads, SP_t, after controlling for global factors. Global factors include the VIX index and returns on the S&P 500 and MSCI Emerging Markets ETF indices. Sample period: January 2007-February 2008.

TABLE 1. Summary Statistics

Moments	Non-Event	Event
Mean $\Delta ln(SP)$	0.055	-0.554
SD $\Delta ln(SP)$	2.717	3.569
Mean ΔBE	0.007	0.029
SD ΔBE	0.189	0.294
Cov($\Delta ln(SP)$,ΔBE)	-0.017	-0.573
Observations	234	24

Notes: The table reports the mean and standard deviation of the daily change in BE_t, the mean and standard deviation of the daily log change in SP_t, and their covariance during event and non-event windows. Event days are defined as 2-day windows around days on which the Argentine government reported the inflation rate. Non-event days are all the others. Sample period: January 2007-February 2008.

days are classified as *non-event days* (NE). The relation between ΔBE_t and $\Delta \ln SP_t$ is not significant for non-event days (blue dots).

Table 1 reports summary statistics on daily changes in the BE rate and sovereign spreads for event and non-event days. The covariance between these variables is close to zero for non-event days but decreases sharply during event days. More importantly, the volatility of ΔBE_t

substantially increases during event days. In the next subsection, we use this difference in volatilities to identify the effect of the misreport on sovereign spreads.

3.3. *Framework and Results*

To allow for the possibility that (i) sovereign spreads may affect ΔBE_t and (ii) there may be unobserved common factors, we consider the following system of equations:

$$\Delta \ln SP_t = \alpha_0 + \alpha_1 \Delta BE_t + \alpha_2 X_t + \epsilon_t \tag{11}$$

$$\Delta BE_t = \beta_0 + \beta_1 \Delta \ln SP_t + \beta_2 X_t + \eta_t, \tag{12}$$

where $\Delta \ln SP_t$ is the log change in sovereign spreads for bonds denominated in dollars and X_t is a vector of common shocks. We further assume that the shocks ϵ_t and η_t have no serial correlation and are uncorrelated with each other and with the common shocks X_t.

Our coefficient of interest is α_1. According to our reputational model, we should expect α_1 to be negative. That is, an increase in the unexpected underreport of inflation (i.e., a decrease in ΔBE_t) should have a negative effect on the government's reputation, leading to a rise in its sovereign spreads.

If we simply run OLS on Equation (11), there are two potential sources of bias: simultaneity and omitted variables. The former appears if $\beta_1 \neq 0$. The latter exists if $\alpha_2 \neq 0$ and $\beta_2 \neq 0$. In order for the OLS estimate of α_1 to be unbiased, an exogenous change in $\Delta \ln SP_t$ must have no effect on ΔBE_t and there must be no omitted common shocks. As previously explained, these two assumptions are implausible in our context.

To tackle these problems, we follow a heteroskedasticity-based identification approach. The formal identifying assumption is that the variance of shocks to ΔBE_t, η_t, is higher around days on which the government announces the inflation rate, while the variances of the common shocks, X_t, and of the shocks to $\Delta \ln SP_t$, ϵ_t, remain invariant. That is,

$$\begin{aligned} \sigma_{\eta,E} &> \sigma_{\eta,NE} \\ \sigma_{\epsilon,E} &= \sigma_{\epsilon,NE} \\ \sigma_{X,E} &= \sigma_{X,NE}. \end{aligned} \tag{13}$$

Let Φ_j be the covariance matrix between $\Delta \ln SP_t$ and ΔBE_t for $j = \{E, NE\}$. If the identifying assumptions hold, it is easy to show that

$$\Delta\Phi = \left(\frac{1}{1-\alpha_1\beta_1}\right)^2 \left[\sigma^2_{\eta,E} - \sigma^2_{\eta,NE}\right] \begin{bmatrix} \alpha_1^2 & \alpha_1 \\ \alpha_1 & 1 \end{bmatrix}, \tag{14}$$

where $\Delta\Phi \equiv \Phi_E - \Phi_{NE}$. From the expression above, it is clear that we can estimate our coefficient of interest in at least two different ways:

$$\hat{\alpha}_1 = \frac{\Delta\Phi_{1,2}}{\Delta\Phi_{2,2}}, \tag{15}$$

$$\tilde{\alpha}_1 = \frac{\Delta\Phi_{1,1}}{\Delta\Phi_{1,2}}. \tag{16}$$

As it is clear from Equation (14), these estimators are relevant only if $\Lambda \equiv \sigma_{\eta,E}/\sigma_{\eta,NE} > 1$. For our identifying assumption to work, the market should thus be surprised about the inflation announced by the government. In Appendix B.7, we show that for the period under analysis (January 2007-February 2008), we can reject the null that $\Lambda = 1$. Interestingly, we cannot reject the null hypothesis during and after the GFC. We interpret this as evidence to suggest that the market learned about the type of government and was no longer surprised by the sequence of misreports.[25]

As shown in Rigobon and Sack (2004), the estimators in Equations (15) and (16) can be implemented in an instrumental-variable framework.[26] Under our null hypothesis, however, $\Delta\Phi_{1,2} = 0$, which renders the $\tilde{\alpha}_1$ estimator inappropriate (see Hébert and Schreger, 2017). For the remainder of the analysis, all of the results are based on the $\hat{\alpha}_1$ estimator.

Table 2 shows the results based on the IV estimator for $\hat{\alpha}_1$. Each column provides the estimates for a different definition of the event and non-event windows. In all of our instrumented regressions, we include a set of global factors to control for aggregate credit market conditions. In particular, we include daily changes in the VIX index, the S&P 500 index, and the MSCI Emerging Markets ETF index. While the addition of these controls is not necessary, given our identifying assumptions, their inclusion allows us to reduce the magnitude of our standard errors.[27]

In all of the specifications, the point estimate $\hat{\alpha}_1$ is negative and statistically significant, which is in line with our reputational channel. Our estimates show that a 1 pp decrease in

[25]It could also be the case that we are capturing a larger volatility in ΔBE during non-event days due to other shocks related to the GFC. However, we cannot reject the null that $\Lambda = 1$ even for the post GFC period (January 2010-February 2011).

[26]We can also compute these estimators directly from the set of moments of Table 1. From the variances and covariances reported in that table, we get $\hat{\alpha}_1 = -10.90$ and $\tilde{\alpha}_1 = -9.64$. Rigobon and Sack (2004) show that the estimators are consistent even if the shocks σ_η, σ_ϵ, or σ_X have heteroskedasticity over time.

[27]Although not shown, all of our results are robust to controlling for changes in the liquidity-premium differential across inflation-linked and dollar-denominated bonds (as proxied by bid-ask-spread differentials).

TABLE 2. Effects of Inflation Misreport on Sovereign Spreads

	(1)	(2)	(3)	(4)
ΔBE	-10.437***	-11.130***	-8.562***	-9.443***
95perc CI	[-15.63, -5.27]	[-17.27, -5.80]	[-13.94, -2.88]	[-14.44, -3.48]
Observations	258	255	67	79
Events	2-day window	3-day window	2-day window	3-day window
Non-events	All other days	All other days	4-day window	4-day window
Controls	Yes	Yes	Yes	Yes

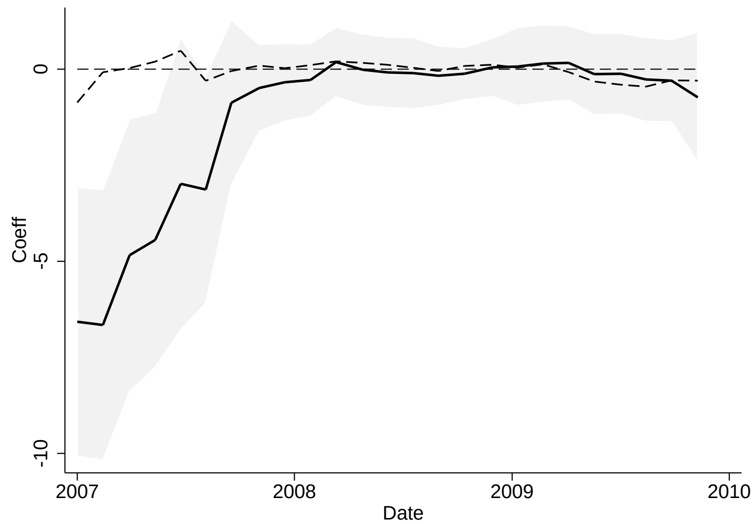
Notes: The table reports the results for the heteroskedasticity-based $\hat{\alpha}_1$ estimator. The dependent variable is $\Delta \ln SP_t$. Definitions of "events" vary across the four columns. Controls include the VIX index, the S&P 500 index, and the MSCI Emerging Markets ETF index. Standard errors and confidence intervals are computed using a stratified bootstrap procedure. 95% confidence intervals are in brackets. ***, **, and * denote significance at 1%, 5%, and 10%, respectively. Sample period: January 2007-February 2008.

ΔBE_t (i.e., an increase in the unexpected underreport of inflation) leads to a 10% increase in sovereign spreads. In terms of economic magnitudes, the reported estimates imply that a 1–sd decrease in ΔBE_t can account for more than two thirds of the daily dispersion of $\Delta \ln SP_t$ (during the event windows).

The results in Table 2 are based on the period between the first misreport and the start of the financial crisis (January 2007-February 2008). For sample periods during and after the GFC, we cannot reject the null hypothesis that $\Lambda = 1$ and we can therefore not apply the Rigobon and Sack methodology. Instead, we consider OLS regressions and a standard event-study analysis, based on 2-day windows around inflation announcements. In these cases, the (stronger) identifying assumption would be that changes in the BE rate during the 2-day windows are driven exclusively by the government's inflation announcement. The estimates are therefore subject to the concern that other factors may have changed during those event days and affected both the BE rate and sovereign spreads. Appendix B.8 provides further details.

Figure 5 presents OLS estimates for α_1 based on rolling windows of 12 months. The black line shows the estimates for α_1 around days on which the government announced the inflation rate. The dashed gray line shows the estimates for all of the other days. The estimates are negative and significant only for event days. Moreover, the estimates are significant only for the first year of the sequence of misreports. Interpreted through the lens of our reputational

Figure 5. OLS Estimates - Rolling Windows



Notes: The figure shows OLS estimates for α_1 based on 12-month rolling windows. The black line shows the estimates around days on which the government announced the inflation rate (event days). The dashed gray line shows the estimates for all of the other days. See Appendix B.8 for details.

model, the results suggest that after 2007, the lenders' prior about the government type (ζ) reached its lower bound, and therefore the misreports no longer affected sovereign spreads.

For the event-study analysis, we classify events as a "good news event" (GNE) or a "bad news event" (BNE) based on the change in BE_t around the government's inflation announcement. Event window j is classified as a BNE if $\mu^{E,j}_{\Delta BE} < \mu^{NE}_{\Delta BE}$, where $\mu^{E,j}_{\Delta BE}$ is the mean daily change in BE_t across event window j, and $\mu^{NE}_{\Delta BE}$ is the mean change across all non-event days in the sample. Appendix B.8 presents the results. For our baseline sample period, the results show an asymmetric response of spreads to news events. In particular, we find a large and positive increase (decrease) in Argentina's sovereign spreads during BNE (GNE). During BNE, for instance, Argentina's sovereign spreads increased on average by 1.5 pp (daily). After March 2008, however, we find no relation between BNE or GNE and changes in Argentina's spreads, which is consistent with the OLS estimates in Figure 5.

3.4. *A Reputational Channel?*

Although the results so far are in line with our reputational channel, other channels may be at play. In this section we consider alternative explanations and provide evidence that supports our reputational mechanism.

A first concern is based on the fact that changes in BE_t may be capturing not only news regarding the misreport but also news about the "true" inflation rate (see, for instance, Nakamura and Steinsson, 2018). If that were the case, news about the true inflation rate could affect the real economy, which in turn ends up affecting sovereign spreads regardless of the Argentine government's reputation. A second concern is that the misreports may induce distortions in the real economy. Regardless of its sign, the misreports could distort relative prices, increase uncertainty, and reduce investment. All of these factors may end up affecting the default risk of the government, regardless of its reputation.

Both of these alternative channels seem at odds with the results presented so far. First, for a high-inflation economy such as Argentina, if $\Delta BE_t < 0$ is actually capturing a lower than expected "true" inflation rate, this may be perceived as a good signal about the fundamentals of the economy, which should reduce the default risk. We should thus expect a positive link between ΔBE_t and ΔSP_t, contrary to the results presented in Table 2. Second, if the misreports are creating distortions in the real economy, we would expect a U-shaped relation between ΔBE_t and $\Delta \ln SP_t$, which is at odds with Figure 4 and our event-study analysis (Appendix B.8), in which we show an asymmetric response of spreads to good news events and bad news events.

To formally address these alternative explanations, we extend our heteroskedasticity-based framework to allow for the possibility that the inflation announcements may affect the fundamentals of the economy directly. To this end, we expand the system of equations in (11)-(12) to account for the potential effects of ΔBE_t on the real economy. We use the daily return (R_t) of an index of publicly traded Argentine firms (MERVAL) to proxy for changes in the real economy. In particular, we consider the following system:

$$\Delta BE_t = \beta_0 + \beta_1 \Delta \ln SP_t + \beta_2 R_t + \beta_3 X_t + \eta_t \tag{17}$$

$$\Delta \ln SP_t = \alpha_0 + \alpha_1 \Delta BE_t + \alpha_2 R_t + \alpha_3 X_t + \epsilon_t \tag{18}$$

$$R_t = \gamma_0 + \gamma_1 \Delta BE_t + \gamma_3 X_t + \nu_t, \tag{19}$$

where we assume that η_t, ϵ_t, ν_t, and X_t are uncorrelated.

In Appendix B.9, we show that even if $\sigma_{\eta,E} > \sigma_{\eta,NE}$, we can no longer identify our parameter of interest, α_1. Under this setup, the heteroskedasticity-based approach allows us to identify γ_1 and $\tilde{\alpha_1} \equiv \alpha_1 + \alpha_2\gamma_1$. Notice that α_1 would account for the *reputational channel*, whereas $\alpha_2\gamma_1$ accounts for a *fundamentals channel*—i.e., the effect of inflation announcements on sovereign spreads through the real economy. Thus, $\gamma_1 \neq 0$ would invalidate our interpretation based on

TABLE 3. Effects of Inflation Misreport on Stock Returns

	(1)	(2)	(3)	(4)
ΔBE	0.246	0.351	0.035	-0.063
95perc CI	[-1.31, 1.71]	[-1.44, 1.91]	[-2.19, 1.41]	[-1.96, 1.33]
Observations	241	238	62	74
Events	2-day window	3-day window	2-day window	3-day window
Non-events	All other days	All other days	4-day window	4-day window
Controls	Yes	Yes	Yes	Yes

Notes: The table reports the results for the heteroskedasticity IV estimator. The dependent variable is R_t. Controls include the VIX index, the S&P 500 index, and the MSCI Emerging Markets ETF index. Standard errors and confidence intervals are computed using a stratified bootstrap procedure. 95% confidence intervals are in brackets. ***, **, and * denote significance at 1%, 5%, and 10%, respectively. Sample period: January 2007-February 2008.

a reputational mechanism. In other words, if $\gamma_1 \neq 0$, the estimates reported in Table 2 may be simply driven by the effects of the inflation announcements on the real economy.

Table 3 presents the IV estimates for γ_1 across different event windows. All of the estimates are small in absolute value, their sign varies with the specification, and none of them are statistically significant. We take this as additional evidence to support our reputational channel. In Appendix B.9, we further extend the system of equations in (17)-(19) to allow for the possibility that the stock market is directly affected by changes in sovereign spreads. This specification is motivated by Hébert and Schreger (2017), who find that increases in sovereign risk lower the returns of the domestic stock market. Under this extended setup, we analytically characterize the biases of our estimates and argue that our estimates for α_1 are downward biased (in terms of magnitudes). Our results, therefore, may be interpreted as a lower bound.

In Appendix B.10, we complement the previous analysis with a structural VAR (as in Mertens and Ravn, 2013; Gertler and Karadi, 2015) that incorporates the interactions between inflation misreports, spreads, and a measure of economic activity. Considering the changes in misreport as a policy variable, we identify structural shocks to the misreport equation using high-frequency changes in the BE inflation rate during event windows. The results are in line with those of Tables 2 and 3. We find that upon a 1–sd structural shock to inflation misreport, spreads increase by 6% on impact. The response of economic activity is lagged and not statistically significant.

4. Quantitative Analysis

In this section we use our empirical estimates, as well as other moments for the Argentine economy, to discipline the reputational model described in Section 2. We then use the calibrated model to back up our model-implied measure of reputation and study the role of fundamentals behind the link between reputation and sovereign spreads. Finally, we use our model to quantify the fraction of Argentina's sovereign spreads that can be attributed to reputation.

4.1. *Calibration and Model Fit*

The model is calibrated at quarterly frequency. The calibration follows a two-step procedure. First, we fix a subset of parameters to values that are either standard in the literature or based on historical Argentine data. We then internally calibrate the remaining parameters to match relevant moments for Argentina's sovereign spreads and other business-cycle statistics. In Appendix C.6, we describe the algorithm used to solve the model.

In terms of functional forms, we assume a CRRA utility function: $u\left(c\right)=\frac{c^{1-\gamma}}{1-\gamma}$, with risk-aversion parameter γ. The endowment follows an AR(1) process given by $\ln\left(y_t\right)=\rho_y\ln\left(y_{t-1}\right)+\epsilon_{y,t}$, with $\epsilon_{y,t}\sim N(0,\sigma_y)$. As in Chatterjee and Eyigungor (2012), the exogenous default cost on income is modeled as $\phi_j\left(y\right)=\max\left\{\left(\bar{\chi}_0+\chi_j\right)y+\bar{\chi}_1 y^2,0\right\}$, where $j=\{C,S\}$, $\bar{\chi}_0<0$, and $\bar{\chi}_1>0$. We set $0>\chi_S=-\chi_C=\bar{\chi}_2$ in order to get a larger default set for the strategic type.

Motivated by the Argentine case, we consider the following specification for the probability of receiving message m. We assume that agents do not observe the inflation misreport $\tilde{\pi}$ but receive a noisy signal about it, $\tilde{\pi}^o$. In particular, we assume that $\tilde{\pi}^o\mid\tilde{\pi}\sim N(\tilde{\pi},\sigma)$, where σ captures the noise of the signal. We further assume that agents detect the misreport (i.e., $m=L$ is realized) if $\tilde{\pi}^o<\alpha$, where $\alpha<0$. Under this setup, $Prob\left(m=L\mid\tilde{\pi}\right)=\Phi_{(\tilde{\pi},\sigma)}\left(\alpha\right)$, where $\Phi_{(\tilde{\pi},\sigma)}$ is the cumulative distribution function of a normal random variable with mean $\tilde{\pi}$ and standard deviation σ. Thus, for a given noise σ, the probability of receiving message $m=L$ is increasing in α.

Table 4 describes the model calibration. Panel (A) lists the parameters we fix in the calibration. We set the risk aversion $\gamma=2$—a standard value in the literature. The real rate is set to $r=1\%$, in line with the observed average real rate in the United States. The reentry parameter is set to $\theta=0.0385$, which implies an average exclusion period from international markets after a default of 6.5 years.[28] We set $\lambda=0.05$ to match an average debt maturity

[28] This measure is taken from Chatterjee and Eyigungor (2012) and is constructed as an average of the time it took Argentina to reach a settlement on the defaulted debt across different default episodes, based on data provided by Beim and Calomiris (2001); Benjamin and Wright (2009); and Gelos et al. (2011).

TABLE 4. Calibration of the Model

Panel A: Fixed Parameters			Panel B: Calibrated Parameters		
Param.	Description	Value	Param.	Description	Value
γ	Risk aversion	2.00	β	Discount rate	0.95
z	Coupon payments	0.03	$\bar{\chi}_0$	Default cost—level	−0.242
λ	Debt maturity	0.05	$\bar{\chi}_1$	Default cost—curvature	0.325
r	Risk-free interest rate	0.01	$\bar{\chi}_2$	Default cost—differential	−0.006
T_{jj}	Persistence j-type	0.969	B	Inflation-indexed debt service	0.02
ρ_y	Endowment, autocorrelation	0.93	α	Probability threshold	−0.028
σ_y	Endowment, shock volatility	0.02			
θ	Reentry probability	0.0385			
σ	Precision of signal	0.011			

of 5 years and $z = 0.03$ to match the debt service, as in Chatterjee and Eyigungor (2012). Regarding the frequency at which the government type changes, we set $T_{CC} = T_{SS} = 0.969$ to reflect an election cycle of 8 years.[29] Parameters for the endowment process, ρ_y and σ_y, are estimated based on log-linearly detrended quarterly real GDP data for Argentina.[30] Lastly, we fix σ to match the noise behind the alternative measures of the Argentine inflation rate (as shown in Figure 3). In particular, we set $\sigma = 0.011$ to match the quarterly cross-sectional volatility across the observed misreports during 2007-2012.

We calibrate the remaining parameters of our model (Panel B of Table 4) to match key data moments of the Argentine economy, detailed in Table 5. In particular, we jointly calibrate the discount factor β and the default cost parameters $\{\bar{\chi}_0, \bar{\chi}_1, \bar{\chi}_2\}$ to target Argentina's average default rate, average external debt, average spread, and volatility of spreads.[31] We target an annual default frequency of 3.3%, since Argentina has defaulted four times since the 1900s.[32]

[29]In Appendix C.3, we analyze the implications of different persistence values.

[30]We use data for the period 1980.Q1-2012.Q4 to compute the log-linear trend for GDP. We allow for a break in the trend in 2002 because Argentina underwent a severe crisis at the end of 2001 that ended with a default. The results are robust to other years and other specifications.

[31]In the model, annualized spreads are given by $SP = \left(\frac{1+r_b(y,b',\zeta')}{1+r}\right)^4 - 1$, where $r_b(y, b', \zeta')$ is the internal rate of return, as implied by $q(y, b', \zeta') = \frac{[\lambda+(1-\lambda)z]}{\lambda+r_b(y,b',\zeta')}$.

[32]Beim and Calomiris (2001) report a default episode in 1956 and another one in 1982. More recently, Argentina defaulted in 2001 and 2014.

Table 5. Targeted Moments

Target	Description	Data	Model
$\mathbb{E}[D/Y]$	Average debt	72%	72%
$\mathbb{E}[SP]$	Average bond spreads	624bp	630bp
$\sigma(SP)$	Volatility spreads	288bp	253bp
$\mathbb{P}[DF]$	Default frequency	3.3%	3.4%
IIB_s/TD_s	Inflation-indexed debt relative service	27%	26%
$\eta_{BE,SP}$	Semi-elasticity BE to spreads	-10.44	-10.32

Notes: The table shows the moments targeted in the calibration and their model counterparts. For data on spreads and debt, the sample period is 1993.Q4-2008.Q1, excluding the default episode that started in December 2001. The semi-elasticity $\eta_{BE,SP}$ is the one computed in Section 3. Model-implied moments are computed based on windows in which the government is not in default.

For the other three moments, we target an average external-debt-to-GDP ratio of 72%, an average spread of 624 basis points (bps), and a standard deviation of spreads of 288 bps.[33,34] We set B to match the share of Argentina's debt services attributed to IIBs between 2007 and 2012 (27%).

Lastly, we internally calibrate the learning parameter α to match the semi-elasticity between Argentina's sovereign spreads and changes in the BE inflation rate described in the empirical analysis of Section 3. To obtain a tight link between model and data, we compute the price for an auxiliary IIB (with the same maturity structure as b) and use that price to compute the BE inflation rate, which is a function of the lenders' prior ζ and the conjectured $\tilde{\Pi}^{\star}_{S}$. Since our empirical elasticity is measured at a high frequency, we extend the baseline model of Section 2 to allow for two instances of trading in secondary markets within a period. The first trading instance (A) is at the beginning of stage 1 and before the message m is realized; the second

[33]These moments are computed for the period 1993.Q4-2008.Q1. Argentina was in default until 1993 and no data for spreads are available prior to that year. From this period, we exclude the 2001.Q3-2005.Q3 subsample because Argentina defaulted in December 2001 and was excluded from debt markets until 2005. We do not include the period of the GFC because our model does not consider mechanisms to explain changes in spreads due to foreign conditions (for instance, changes in risk aversion).

[34]As in Chatterjee and Eyigungor (2012), we match only a portion of Argentina's external debt because we do not model repayment. In Argentina's case, the repayment of defaulted debt has been around 30%.

(B) occurs after lenders observe message m and update their beliefs (i.e., $\hat{\zeta}(m)$) accordingly. See Appendix A.4 for further details.

Under this extension, we can capture changes in the BE inflation rate and spreads induced by an update in lenders' beliefs about the government type (i.e., reputation) coming from the realization of message m. Our timing assumption (as described in Figure 1) implies that b' is chosen before the trading instance A. Thus, changes in the BE rate and spreads between trading instances A and B only capture the information provided by the message m and are not driven by changes in the bond policy.

We denote $\Delta BE\left(m\right) \equiv \Delta BE\left(y, b, \tilde{\zeta}, \hat{\zeta}(m)\right)$ and $\Delta \ln SP\left(m\right) \equiv \Delta \ln SP\left(y, b, \tilde{\zeta}, \hat{\zeta}(m)\right)$ to be the changes in prices between trading instances A and B, conditional on the realized message m. Because both $\Delta BE\left(m\right)$ and $\Delta \ln SP\left(m\right)$ are endogenous variables, in order to isolate the causal effect of the misreport on spreads, we construct a counterfactual in which we shock the optimal misreport policy by $\epsilon_{\tilde{\pi}}$. This shock affects the realization of message m and hence the posterior and bond prices. Let m be the realized message under the optimal policy, $\tilde{\pi}^{\star} \equiv \tilde{\pi}^{\star}\left(y, b, \tilde{\zeta}\right)$, and let m_{ϵ} be the realized message under the counterfactual in which the misreport is $\tilde{\pi}^{\star} + \epsilon_{\tilde{\pi}}$. Our model-implied elasticity is defined as

$$\eta_{BE,SP} \equiv \mathbb{E}\left[\frac{\Delta \ln SP\left(m_{\epsilon}\right) - \Delta \ln SP\left(m\right)}{\Delta BE\left(m_{\epsilon}\right) - \Delta BE\left(m\right)}\right]. \tag{20}$$

The last row of Table 5 shows that the model is able to match our empirical elasticity. In Appendix C.2, we describe how changes in α affect $\eta_{BE,SP}$ and provide a sensitivity analysis.

We next assess how the model performs in terms of untargeted moments for both standard and model-specific ones. Table 6 shows that our calibrated model captures key business-cycle moments of the Argentine economy. In particular, it closely approximates the relative volatility and correlation of consumption and trade balance with output. The model also captures the negative correlation between spreads and output, which is a common feature of emerging economies (see, for example, Neumeyer and Perri, 2005 and Aguiar and Gopinath, 2007).

Table 7 shows a set of untargeted moments that are specific to our model. The top panel shows that the model is roughly consistent with the average quarterly misreport, its volatility, and the negative relation between the misreport and the output cycle. For the data counterparts, although we do not observe the actual misreport of inflation, we proxy it by computing the difference between the inflation announced by the government and the average across alternative measures of inflation. The bottom panel compares different moments based on high-frequency changes around days on which the government announces the inflation rate. In the model, we

Table 6. Untargeted Moments: Business-cycle Statistics

Target	Description	Data	Model
$\sigma(\log C)/\sigma(\log Y)$	Relative volatility consumption	1.13	1.30
$\sigma(TB/Y)/\sigma(\log Y)$	Relative volatility trade balance	0.32	0.44
$\text{corr}(\log C, \log Y)$	Correlation consumption & endowment	95%	96%
$\text{corr}(TB/Y, \log Y)$	Correlation trade balance & endowment	−31%	−50%
$\text{corr}(SP, \log Y)$	Correlation spreads & endowment	−42%	−70%

Notes: The table compares a set of untargeted data moments with their model counterparts. Data for consumption and output are for the period 1980-2012 and exclude the 1982 and 2001 default episodes. Data for spreads and trade balance start in 1993. The terms $\log C$ and $\log Y$ denote the log-linear cycle for consumption and output, respectively.

Table 7. Untargeted Moments: Misreport, BE, and Spreads

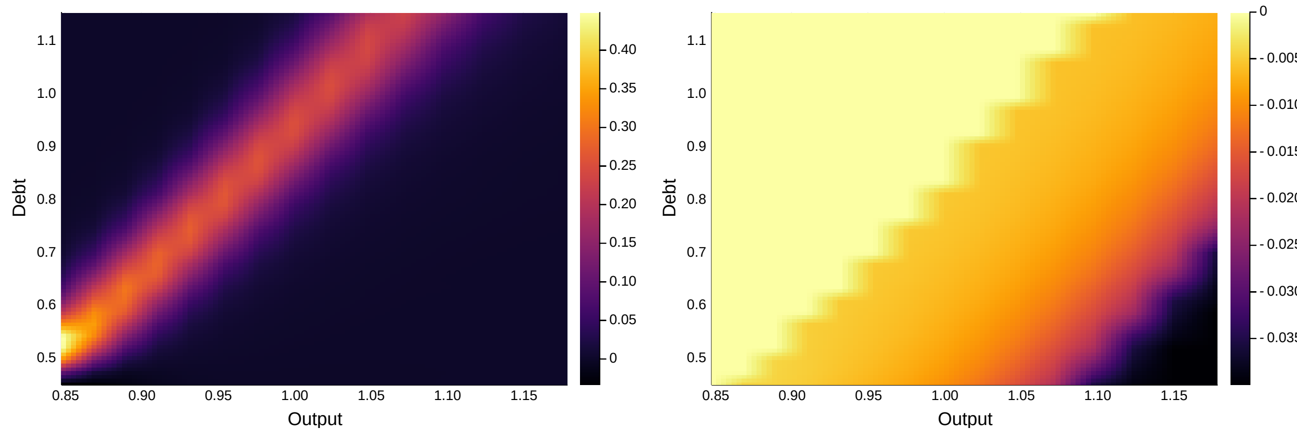
Target	Description	Data	Model
Panel A: Quarterly Frequency			
$\mathbb{E}[\tilde{\pi}]$	Average inflation misreport	−3.47%	−1.82%
$\sigma(\tilde{\pi})$	Volatility inflation misreport	2.31%	0.89%
$\text{corr}(\tilde{\pi}, \log Y)$	Correlation misreport & output	−58%	−31%
Panel B: High Frequency			
$\sigma(\Delta BE)$	Volatility break-even inflation	0.29%	0.15%
$\text{corr}(\epsilon_{\tilde{\pi}}, \Delta BE)$	Correlation misreport & break-even inflation	32%	40%
$\text{corr}(\epsilon_{\tilde{\pi}}, \Delta \ln SP)$	Correlation misreport & spreads	−37%	−39%

Notes: The table compares a set of moments that are specific to our model with their data counterpart. The last three rows show high-frequency changes around days on which the government announces the inflation rate. In the model, we compute these changes by comparing prices at trading instances A and B (i.e., before and after the realization of message m). For the data column, ΔBE and $\Delta \ln SP$ are computed for event days only, as described in Table 1, and are based on the period 2007-2008.

compute these changes by comparing bond prices before and after the realization of message m_ϵ, based on an unexpected shock to the misreport $\epsilon_{\tilde{\pi}}$. In the data, although unobservable, we proxy $\epsilon_{\tilde{\pi}}$ as the change in the observed misreport across two consecutive months; see Appendix B.5 for the rationale. The model matches the volatility of ΔBE and its correlation with $\epsilon_{\tilde{\pi}}$. It

Figure 6. Fundamentals and Reputation

(A) $q(y, b, \zeta_H) - q(y, b, \zeta_L)$ (B) $\tilde{\pi}(y, b, \zeta_H)$



Notes: Panel (A) shows the effect of a change in a government's reputation (from ζ_L to ζ_H) on its bond prices for different combinations of (y, b). Panel (B) shows the optimal $\tilde{\pi}$ policy for different states (y, b). In both cases, the upper left area coincides with points in the state space in which both the C- and S-type default.

also captures the negative correlation between $\epsilon_{\tilde{\pi}}$ and $\Delta \ln SP$ that we observe in the Argentine data.

4.2. *Links between Reputation and Fundamentals*

In what follows, we use the model to disentangle the fraction of a government's spreads that can be explained by its reputation. We refer to this measure as the "reputation premium." We then analyze the role of macroeconomic fundamentals (y and b) behind the link between reputation and spreads.

We first analyze the effect of reputation on bond prices and describe the optimal $\tilde{\pi}$ policy for different points of the state space. Panel (A) of Figure 6 shows the effect of a change in ζ on bond prices for different combinations of (y, b). On the upper left corner (high b, low y), both the C- and S-type choose to default, and therefore the effect of ζ on bond prices is negligible. On the lower right corner, on the other hand, the default probability for the C- and S-type is close to zero and changes in reputation do not affect bond prices. On the main diagonal, however, the default probability of the S-type is (weakly) larger than that of the C-type; see Appendix Figure C.1. Therefore, in these points of the state space, changes in ζ do affect bond prices significantly. Panel (B) shows that as we approach this area, the S-type optimally chooses to decrease the magnitude of $\tilde{\pi}$, since it does not want to reveal its type.

Table 8. Decomposition of Spreads: The Reputation Premium

Moment	Description	Value
$\mathbb{E}[\Upsilon]$	Average reputation premium	98bp
$\mathbb{E}[\Upsilon/SP]$	Incidence reputation premium on spreads	13%
$\sigma(\Upsilon)/\sigma(SP)$	Reputation-premium volatility	44%
$\sigma(SP\|\zeta_H)/\sigma(SP)$	Spread volatility under high reputation	60%
$\mathbb{E}[\Upsilon/SP\|Y < Y_l]$	Incidence with low output	21%
$\text{corr}(\Upsilon, \log Y)$	Correlation reputation premium & output	−64%
$\text{corr}(\Upsilon/SP, \log Y)$	Correlation reputation incidence & output	−67%

Notes: The table reports moments related to the reputation premium, Υ, and the link between Υ and the economy's fundamentals.

We define *reputation premium* (Υ) as the additional borrowing cost a government faces for not having a high reputation. Formally, it is the difference between realized (i.e., observed) sovereign spreads and those under a counterfactual in which the government's reputation is the maximum possible. That is,

$$\Upsilon(y, b, \zeta) \equiv SP(y, b, \zeta) - SP(y, b, \zeta_H), \tag{21}$$

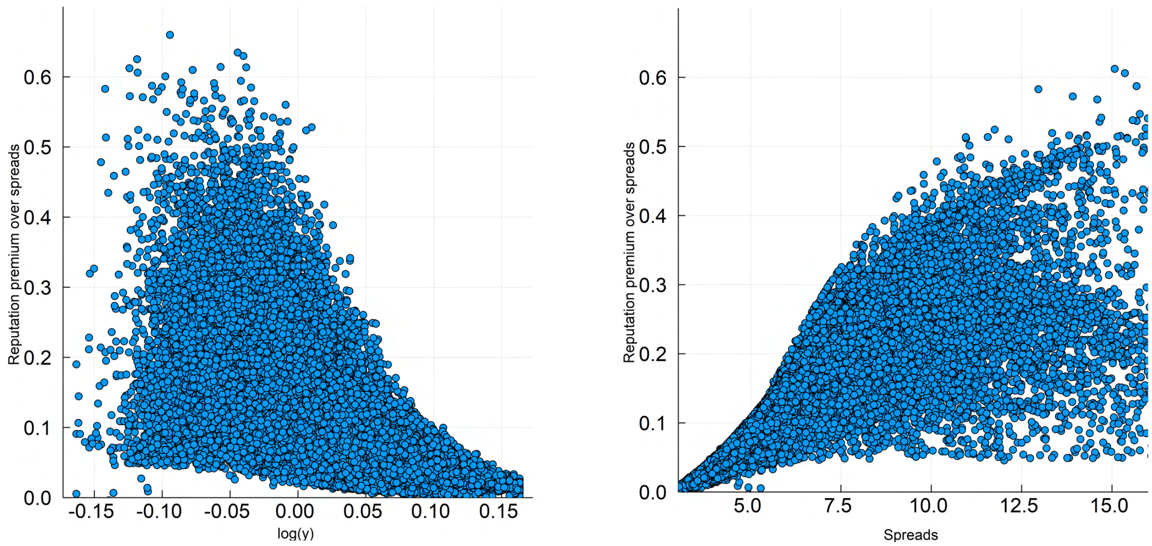
where ζ_H is the upper bound for the lenders' prior.

Table 8 provides different moments describing the reputation premium based on model simulations. On average, the premium is 98 bps, which accounts for 13% of sovereign spreads. More importantly, the table shows that spreads in our model would be 40% less volatile absent the reputation premium.

The bottom panel of Table 8 shows that the incidence of the reputation premium is state contingent: (i) the premium accounts for about 20% of spreads when output is one standard deviation below its average, and (ii) the correlation between reputation premium and output is highly negative. Figure 7 illustrates these points in more detail. It shows the fraction of spreads explained by the reputation premium for different values of output (Panel A) and observed spreads (Panel B). The figure shows that the reputation premium can account for up to 50% of spreads when the economy is in a severe recession or when the default probability is high.

FIGURE 7. Reputation Premium and Fundamentals

(A) Reputation Premium vs Output (B) Reputation Premium vs Spreads



Notes: The figure shows the incidence of the reputation premium for different macroeconomic fundamentals. Panel (A) shows the reputation of premium (as a share of spreads) for different levels of output. Panel (B) shows the same measure but for different levels of observed spreads.

4.3. *The Costs of Information Frictions*

We next analyze the costs of information frictions about the government type. In particular, we study, from the perspective of the C-type, the spreads that it faces in our baseline model and under two counterfactuals. The first one (*fixed C-type* case) is a counterfactual in which the type of government is fixed (and known by lenders). The second one (*perfect-information* case) is a counterfactual in which the type of government is time varying but lenders can perfectly observe it. Table 9 presents the results.

In the baseline model, even after we condition for periods in which the C-type is in charge of the government, the reputation premium still accounts for a sizable share of spreads and explains a third of its volatility. This is because lenders' beliefs are slow moving and the C-type cannot perfectly reveal its type to lenders. Under the *fixed C-type case*, the government attains a much larger debt and, at the same time, faces lower spreads. This result thus highlights that the mere existence of the S-type significantly affects the borrowing costs faced by the C-type. Our model can thus help explain why countries in which political parties have displayed different preferences over default in the past may face higher spreads.

Table 9. The Costs of Information Frictions

	Debt and Spreads			Reputation Premium	
	$\mathbb{E}[D/Y]$	$\mathbb{E}[SP]$	$\sigma(SP)$	$\mathbb{E}[\Upsilon]$	$\sigma[\Upsilon]$
Baseline Model	73%	653bp	267bp	92bp	93bp
Fixed C-type	83%	497bp	202bp	-	-
Perfect Information	76%	593bp	235bp	-	-

Notes: The table shows moments for debt and spreads, conditioning on cases in which the government is of the C-type. The first row shows the results for our baseline model with alternating types and imperfect information. The second row shows a counterfactual in which the type of government is fixed (and observable). The last row shows the case in which the type of government alternates but it is perfectly observable.

The last row of Table 9 provides a counterfactual in which the type of government is perfectly observable, which allows us to disentangle the implications of information frictions. Under perfect information, the C-type avoids paying the reputation premium and is thus able to increase its stock of debt and, at the same time, face lower spreads.[35] While the C-type is negatively affected by not being able to perfectly reveal its type, the S-type may benefit from it, so the overall effects of information frictions in terms of welfare are not clear.[36] In Appendix C.4, we provide a welfare analysis and show that on average, the government would be better off in the perfect-information case.

4.4. *The Argentine Case*

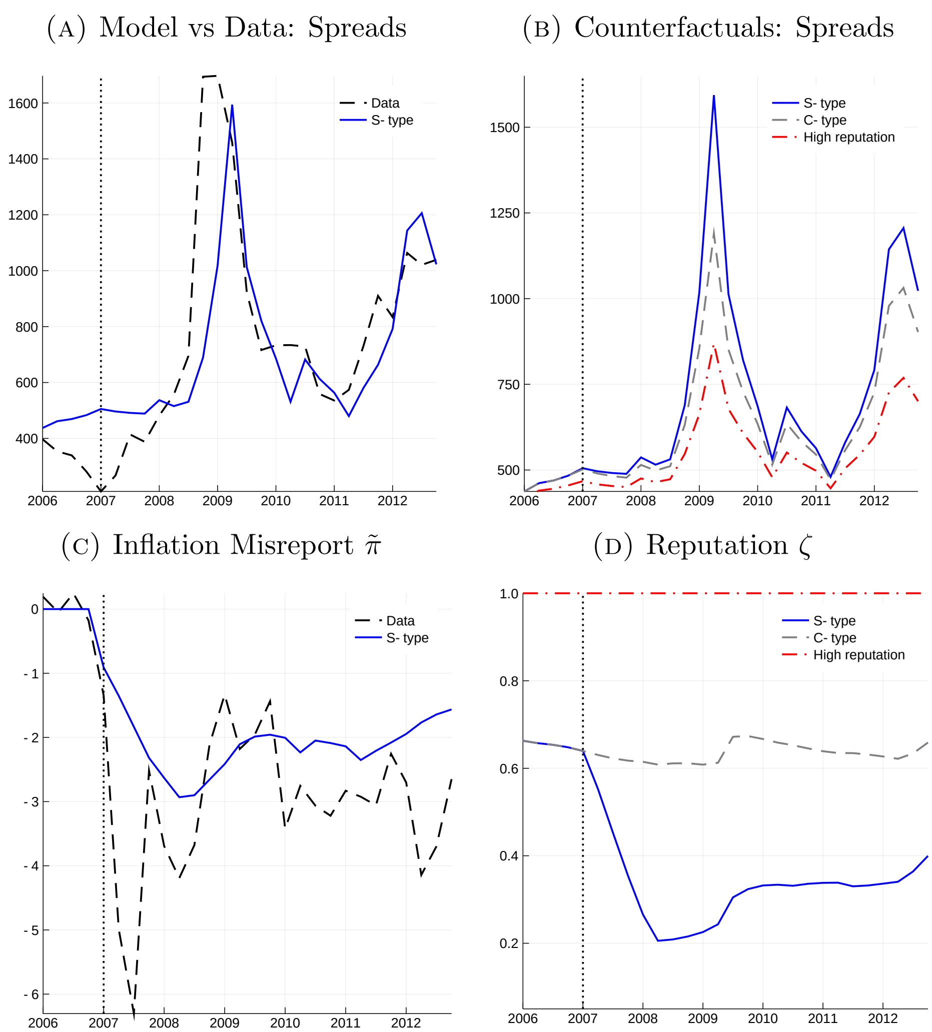
We use the calibrated model to simulate Argentine spreads during the period 2006.Q1-2012.Q4. To this end, we feed the model with the observed evolution of Argentina's log-linear cycle of GDP; see Appendix Figure C.8a for the path for output. We choose the initial value of debt to match the observed spreads in 2006.Q1. We assume that the government is initially of the C-type and becomes of the S-type starting in 2007.Q1.[37] We simulate the economy 20,000 times and take averages across simulations. Starting in 2007.Q1, each simulation i differs in its realized sequence of messages, $\{m_t^i\}_{t=1}^T$. We filter by simulations in which at least one message $m = L$ was realized during the baseline period of our empirical analysis in Section 3 (i.e.,

[35]Differences in spreads across the counterfactuals are even more striking once we fix the level of debt. See Appendix C.4.

[36]Although not reported in Table 9, conditional on periods in which the government is of the S-type, the average spread is around 600 bps under imperfect information and 615 bps under perfect information.

[37]We set the initial reputation to match the average reputation under the C-type.

Figure 8. Model Simulations: Comparison with Data and Counterfactuals



(A) Model vs Data: Spreads (B) Counterfactuals: Spreads

(C) Inflation Misreport $\tilde{\pi}$ (D) Reputation ζ

Notes: Panel (A) shows the evolution of Argentina's sovereign spreads for the period 2006.Q1-2012.Q4 (dashed black line) and model-implied dynamics. The solid blue line shows the average spread across the simulations. Panel (B) shows the model-implied dynamics under the S-type scenario (solid blue line), the C-type scenario (dashed gray line), and a high-reputation counterfactual (dashed red line). Panel (C) shows the optimal misreport policy $\tilde{\pi}$. Panel (D) shows the evolution of ζ under the different counterfactuals.

2007.Q1-2008.Q1). In Appendix C.5, we compare the time-series implications of our model with the perfect information counterfactual.

Panel (A) of Figure 8 shows the dynamics of spreads in the data (dashed line) and the model-implied dynamics (solid blue line). The model provides a path for Argentine spreads that moves in line with that of the data. In particular, it is able to account for a large fraction of the increase in Argentina's spreads during the GFC, even though we consider risk-neutral lenders and abstract from changes in risk premia.[38]

The other panels of Figure 8 compare the implied dynamics of our baseline simulation against two counterfactuals. In the first one, the government remains of the C-type and its reputation varies based on the realized messages and the Markov chain T. In the second counterfactual, we assume that the reputation is fixed at ζ_H. For both counterfactuals, we set the same path of debt as that observed under the baseline simulation.[39]

Starting in 2007.Q1, Panel (B) shows that spreads under the baseline simulation (solid blue line) start to decouple from those implied by the counterfactuals (gray and red lines). This is because the S-type reveals its type by setting $\tilde{\pi} < 0$ (Panel C), which implies that message $m = L$ is realized more frequently and the government's reputation declines (Panel D). Even though it is untargeted in the calibration, the model is able to match the observed paths of misreports reasonably well.

Overall, the previous analysis highlights that the decrease in reputation led to a striking additional response of Argentina's spreads. In particular, Argentina's loss of reputation can explain 30%-50% of the increase in its sovereign spreads during the GFC.

5. Conclusion

In this paper, we study how a government's reputation is shaped by its policies and quantify how markets price this reputation. To this end, we focus on a debt-repayment setting in which reputation is a first-order concern. We develop a sovereign default model with uncertainty about the government type and noisy signals. In the model, agents observe signals about the government's policies and use those signals to update their beliefs about its type (i.e., reputation). Changes in reputation affect the lenders' perceived probability of default and therefore sovereign spreads. Guided by the model, we use the 2007-2012 Argentine episode of inflation misreport to provide new empirical evidence on the link between reputation and borrowing costs. We argue that this policy provided (noisy) information to lenders regarding the type of government, which affected its reputation. We find that the market priced the

[38]For papers that study the role of global factors and international lenders, see, for example, Borri and Verdelhan (2011); Lizarazo (2013); Aguiar et al. (2016); Bai, Perri, and Kehoe (2019); Bocola and Dovis (2019); Morelli, Ottonello, and Perez (2022).

[39]This way we can isolate changes in spreads that are simply driven by differences in the stock of debt.

sequence of misreports, as reflected in a significant increase in the spreads of Argentina's dollar-denominated bonds. Our quantitative model shows that changes in reputation can have long-lasting effects. In particular, we find that the loss in Argentina's reputation due to the misreport is crucial for matching the observed excess sensitivity of Argentina's spreads during the GFC and, to some extent, its posterior decoupling from the rest of the region.

More generally, our results stress the role of reputation as a type of gained capital that is salient for policymakers. Reputation and the existence of asymmetric information can affect other areas of policy interest, such as the effectiveness of government stabilization policies, the rule of law and a country's investment environment, international trade and relations with foreign countries and organizations, and government contracts with other entities. We leave a more detailed analysis of the role of reputation in these areas to future research.

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Appendix A. A Reputational Model of Sovereign Default

In this appendix, we provide the details for the model described in Section 2. We start by describing the government's recursive problem and the bond-pricing kernel. We then define the equilibrium for this economy. Lastly, we provide an extension of the model that includes a secondary market for government bonds. This extension allows us to compute a model-implied semi-elasticity between changes in the BE inflation rate and sovereign spreads, which we can link to our empirical analysis.

A.1. *Government's Recursive Problem*

We outline the j-type government's recursive problem. We first describe the optimal default decision. We then describe the optimal debt and $\tilde{\pi}$ policies.

Stage 0: Optimal Default Decision

At stage 0, assuming the country is currently out of default, the government chooses whether to default or not. Each type j solves the following problem:

$$W_j\left(y,b,\zeta\right) = \max_{d\in\{0,1\}} \left\{ W_j^R\left(y,b,\tilde{\zeta}\right),\, W_j^D\left(y,\tilde{\zeta}\right)\right\}, \tag{A.1}$$

where $W_j^R(\cdot)$ denotes the value function in case of repayment, $W_j^D(\cdot)$ is the value function in case of default, and $\tilde{\zeta} \equiv \tilde{\zeta}\left(d,\zeta;d_S^{\star},d_C^{\star}\right)$ denotes the lenders' posterior (defined in Equation (3)), which depends on the lenders' conjectures $(d_S^{\star}, d_C^{\star})$. Let $d_j(y,b,\zeta) = \{0,1\}$ denote the optimal default policy for type j.

If the government defaults, it faces an output cost $\phi_j(y)$ and is temporarily excluded from international debt markets. We assume that it regains access to debt markets with probability θ. There is no recovery value and the stock of debt is $b = 0$ after exiting a default. The value of default for the j-type is given by

$$\begin{aligned} W_j^D\left(y,\tilde{\zeta}\right) &= u\left(y-\phi_j\left(y\right)\right) + \\ &+\theta\beta \int_y \left\{ T_{jj} W_j\left(y',0,\zeta'\right) + T_{j(-j)} W_{(-j)}\left(y',0,\zeta'\right)\right\} dF\left(y' \mid y\right) \\ &+\left[1-\theta\right]\beta \int_y \left\{ T_{jj} \tilde{W}_j^D\left(y',\zeta'\right) + T_{j(-j)} \tilde{W}_{(-j)}^D\left(y',\zeta'\right)\right\} dF\left(y' \mid y\right) \\ s.t. \quad \zeta' &= T_{CC}\times\tilde{\zeta} + T_{SC}\times\left(1-\tilde{\zeta}\right), \end{aligned} \tag{A.2}$$

where $(-j)$ refers to the type other than j and $\tilde{W}_j^D(\cdot)$ denotes the value function if the government is already in default. It is given by

$$\begin{aligned} \tilde{W}_j^D(y,\zeta) = & \, u(y-\phi_j(y)) + \\ & +\theta\beta \int_y \left\{ T_{jj} W_j(y',0,\zeta') + T_{j(-j)} W_{(-j)}(y',0,\zeta') \right\} dF(y' \mid y) \\ & + [1-\theta]\beta \int_y \left\{ T_{jj} \tilde{W}_j^D(y',\zeta') + T_{j(-j)} \tilde{W}_{(-j)}^D(y',\zeta') \right\} dF(y' \mid y) \\ s.t. \quad & \zeta' = T_{CC} \times \zeta + T_{SC} \times (1-\zeta). \end{aligned} \tag{A.3}$$

Notice that the only difference between Equations (A.2) and (A.3) is the evolution of the posterior.

Stage 1: Optimal Debt Issuance and $\tilde{\pi}$ Policies

At the beginning of stage 1, lenders have adjusted their beliefs based on the observed choice of d. The economy's state is thus given by $(y, b, \tilde{\zeta})$. If the government is not in default, it then chooses its optimal debt issuance and $\tilde{\pi}$ policies.

Since the goal of the analysis is to focus on the information provided by the $\tilde{\pi}$ policy, we assume that bond policies are uninformative about the type of government. To this end, as in Amador and Phelan (2021), we assume that both the C- and S- type follow the same debt policy $b^{\star\prime}\left(y, b, \tilde{\zeta}\right)$.[40] We interpret $b^{\star\prime}(\cdot)$ as a fiscal rule that is not under the control of the j-type. Instead of imposing an arbitrary fiscal rule, we assume that bond policies are optimally chosen by another agent of the economy: the Congress. We assume that the Congress does not observe the type of government and has the same information set as that of the market. Thus, the priors and conjectures of lenders and Congress are the same, which implies that the bond policy is completely uninformative about the government type.

Under these assumptions, given the state $\left(y, b, \tilde{\zeta}\right)$ and the conjectured $\tilde{\pi}$ policies, the bond policy rule $b^{\star\prime} \equiv b'(y, b, \tilde{\zeta})$ is obtained from the following problem:

$$\begin{aligned} b^{\star\prime} = \text{ArgMax} \quad & \tilde{\zeta} \left\{ \sum_{M=\{L,NL\}} Prob\left(m = M \mid \tilde{\Pi}_C^\star\right) V_C\left(\tilde{\Pi}_C^\star, y, b, \hat{\zeta}(M)\right) \right\} + \\ & \left(1-\tilde{\zeta}\right) \left\{ \sum_{M=\{L,NL\}} Prob\left(m = M \mid \tilde{\Pi}_S^\star\right) V_S\left(\tilde{\Pi}_S^\star, y, b, \hat{\zeta}(M)\right) \right\}, \end{aligned} \tag{A.4}$$

[40] In a continuous-time infinite-horizon model with perfectly observed actions and no exogenous cost of default, Amador and Phelan (2021) show that this restriction is without loss of generality. This is because the S-type does not have incentives to completely reveal its type by choosing a different bond policy.

where $\hat{\zeta}(m) \equiv \hat{\zeta}\left(m, \tilde{\zeta}; \tilde{\Pi}_S^*, \tilde{\Pi}_C^*\right)$ is given by Equation (4). The function $V_j(\cdot)$ is defined as[41]

$$V_j\left(\tilde{\pi}, y, b, \hat{\zeta}(m)\right) = u(c) + \beta \int_y \left\{T_{jj} W_j(y', b^{\star\prime}, \zeta') + T_{j(-j)} W_{(-j)}(y', b^{\star\prime}, \zeta')\right\} dF(y' \mid y), \tag{A.5}$$

where $c = y - b[\lambda + (1-\lambda)z] + q(y, b^{\star\prime}, \zeta')[b^{\star\prime} - (1-\lambda)b] + \Omega(\tilde{\pi})$ and ζ' is given by Equation (5). When solving for the optimal bond policy, the Congress takes the pricing kernel $q(y, b', \zeta')$ as given but internalizes the effects of a larger bond issuance on bond prices.

Regarding the optimal $\tilde{\pi}$ policy, taking as given the bond policy $b^{\star\prime}$, the S-type solves the following problem:

$$\begin{aligned} W_S^R\left(y, b, \tilde{\zeta}\right) = & \max_{\tilde{\pi}} \sum_{M=\{L,NL\}} Prob\left(m = M|\tilde{\pi}\right) \times V_S\left(\tilde{\pi}, y, b, \hat{\zeta}(M)\right) \\ & s.t. \quad \tilde{\pi} \in [\underline{\pi}, 0]. \end{aligned} \tag{A.6}$$

Figure A.1 provides a graphical illustration behind the optimal choice of $\tilde{\pi}$. For a given realization of message m, under the assumption that $\Omega(.)$ is increasing in the magnitude of $\tilde{\pi}$, then $V_S\left(\tilde{\pi}, y, b, \hat{\zeta}(m)\right)$ is increasing in $\mid \tilde{\pi} \mid$. It also attains a higher value when $m = NL$ because of the effect the message has on reputation and bond prices. The dashed line depicts the weighted average between $V_S\left(\tilde{\pi}, y, b, \hat{\zeta}(NL)\right)$ and $V_S\left(\tilde{\pi}, y, b, \hat{\zeta}(L)\right)$, where the weights are given by $Prob(m \mid \tilde{\pi})$. When choosing $\tilde{\pi}$ the government internalizes its effects on the probability that message m is being realized. Therefore, the choice of $\tilde{\pi}$ involves a stochastic trade-off between higher current consumption and lower reputation. We denote $\tilde{\pi}_S(y, b, \tilde{\zeta})$ to be the optimal policy for the strategic type. As for the C-type, since it commits to $\tilde{\pi} = 0$, we can define its value function at stage 1 as

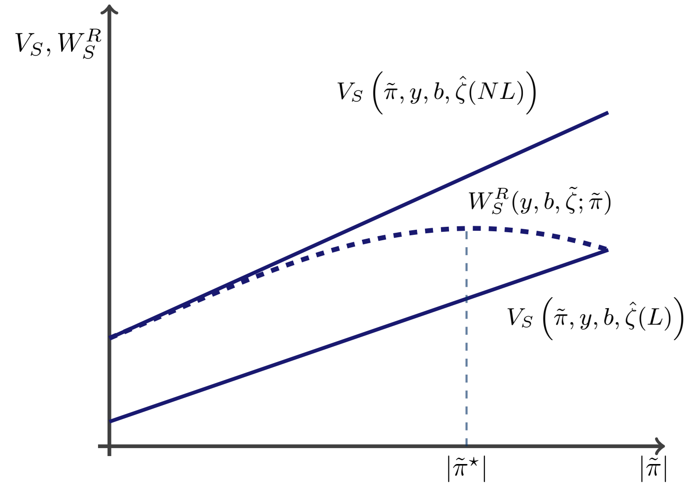
$$W_C^R\left(y, b, \tilde{\zeta}\right) = \sum_{M=\{L,NL\}} Prob\left(m = M|\tilde{\pi} = 0\right) \times V_C\left(0, y, b, \hat{\zeta}(M)\right). \tag{A.7}$$

Stage 2: Bond Issuances in the Primary Market

At stage 2, after message m has been realized, the primary market for bonds opens and the government issues $b^{\star\prime}$. Our timing assumption implies that the Congress does not adjust its choice of b' based on the realization of m. As we explain in Appendix A.4, this assumption allows us to isolate the effect of the message m on spreads (i.e., the reputational effect), which is important to match the empirical semi-elasticity of our empirical analysis.

[41]Notice that $b^{\star\prime}$ is also an argument of $V_j(\cdot)$ but we omit it for notation easiness.

FIGURE A.1. Optimal Choice of $\tilde{\pi}$ - Graphical Illustration



Notes: The figure shows the value function $V_S(\cdot)$ as a function of $\tilde{\pi}$ for the two possible realizations of the message m. The dashed line depicts the linear combination between the two value functions, where the weights depend on the probability of message m being realized, given $\tilde{\pi}$.

A.2. *Pricing Kernels*

We assume that bonds are priced by risk-neutral investors. Let r denote the risk-free rate at which they discount payoffs. Let ζ' be the updated end-of-period posterior as defined in Equations (3)-(5). Let $VR_j(y', b', \zeta')$ be the next-period value of repayment if the government is of the j-type. The bond-pricing kernel is

$$q(y, b', \zeta') = \frac{1}{1+r}\int_y \left\{\zeta' VR_C(y', b', \zeta') + (1-\zeta') VR_S(y', b', \zeta')\right\} dF(y' \mid y) \tag{A.8}$$

with

$$VR_j(y', b', \zeta') \equiv \left(1 - d_j^{\star\prime}\right) \times \left[\sum_{M=\{L,NL\}} Prob(m' = M \mid \tilde{\Pi}_j^{\star\prime})\left(\lambda + (1-\lambda)\left[z + q'_M\right]\right)\right], \tag{A.9}$$

where $d_j^{\star\prime} \equiv d_j^\star(y', b', \zeta')$ refers to the (conjectured) next-period default choice for type j, given the next-period initial state. Similarly, $\tilde{\Pi}_j^{\star\prime} \equiv \tilde{\Pi}_j^\star(y', b', \tilde{\zeta}')$ refers to the conjectured next-period optimal $\tilde{\pi}$ policy, with $\tilde{\zeta}' \equiv \tilde{\zeta}\left(d' = 0, \zeta'; d_S^{\star\prime}, d_C^{\star\prime}\right)$ (as defined in Equation (3)). The term q'_M refers to the next-period price for one unit of debt. This price is contingent on the realization of message m' and is given by

$$
\begin{aligned}
q'_M &= q\left(y', b'', \zeta''\right) \\
\hat{\zeta}' &= \hat{\zeta}\left(M, \tilde{\zeta}'; \tilde{\Pi}^{\star\prime}_S, \tilde{\Pi}^{\star\prime}_C\right) \text{ [as defined in Equation (4)]} \\
\zeta'' &= T_{CC} \times \hat{\zeta}' + T_{SC} \times \left(1 - \hat{\zeta}'\right) \\
b'' &= b^{\star\prime}\left(y', b', \tilde{\zeta}'\right) \text{ [as defined in Equation (A.4)].}
\end{aligned}
$$

A.3. *Definition of Equilibrium*

Definition **1.** *Perfect Bayesian Equilibrium (PBE)*

A PBE is a collection of value functions $\left\{W_j(\cdot), W_j^R(\cdot), W_j^D(\cdot), \tilde{W}_j^D(\cdot), V_j(\cdot)\right\}_{j=\{C,S\}}$*; policy functions* $\{d_j(\cdot), \tilde{\pi}_j(\cdot), b'(\cdot)\}_{j=\{C,S\}}$*; lenders' conjectures* $\left\{d_j^\star(\cdot), \tilde{\Pi}_j^*(\cdot)\right\}_{j=\{C,S\}}$*; lenders' system of beliefs* $\left\{\tilde{\zeta}(\cdot), \hat{\zeta}(\cdot), \zeta'(\cdot)\right\}$*; and bond prices* $q(\cdot)$ *such that:*

1. *Given* $(d_S^\star(\cdot), d_C^\star(\cdot))$*, the posterior* $\tilde{\zeta}(d, \zeta; d_S^\star, d_C^\star)$ *is derived from Equation (3).*
2. *Given* $\left(\tilde{\Pi}_S^\star(\cdot), \tilde{\Pi}_C^\star(\cdot)\right)$*, the posterior* $\hat{\zeta}\left(m, \tilde{\zeta}, \tilde{\Pi}_S^\star, \tilde{\Pi}_C^\star\right)$ *is derived from Equation (4) and* $\zeta'\left(\hat{\zeta}\right)$ *is obtained from Equation (5).*
3. $b'(\cdot)$ *solves the problem in Equation (A.4), where* $V_j(\cdot)$*, as defined in Equation (A.5), is the associated value function.*
4. *Given the value function* $V_S(\cdot)$*,* $\tilde{\pi}_S(\cdot)$ *solves the problem in Equation (A.6) and* $W_S^R(\cdot)$ *is the associated value function. As for the C-type,* $\tilde{\pi}_C(\cdot) = 0$ *(by assumption) and* $W_C^R(\cdot)$*, defined in Equation (A.7), is the associated value function.*
5. *The value functions in case of default,* $W_j^D(\cdot)$ *and* $\tilde{W}_j^D(\cdot)$*, are consistent with Equations (A.2) and (A.3).*
6. *Given the value functions* $W_j^R(\cdot)$ *and* $W_j^D(\cdot)$*,* $d_j(\cdot)$ *solves the problem in Equation (A.1) and* $W_j(\cdot)$ *is the associated value function.*
7. *Given lenders' conjectures* $d_j^\star(\cdot)$ *and* $\tilde{\Pi}_j^\star(\cdot)$*, bond prices are consistent with Equations (A.8) and (A.9).*
8. *Lenders' conjectures coincide with optimal policies:* $d_j^\star(\cdot) = d_j(\cdot)$*,* $\tilde{\Pi}_j^\star(\cdot) = \tilde{\pi}_j(\cdot)$*.*

A.4. *Model Extension: Secondary Markets and Link with Empirical Analysis*

Our empirical semi-elasticity between the BE inflation rate and sovereign spreads (as computed in Section 3) relies on high-frequency data. In particular, it is constructed in a short

Figure A.2. Timing of Events: Infinite-period Model

	If default	If no default	
Stage 0	Stage 1	Stage 1	Stage 2
- Initial $\mathbf{S} = (y, b, \zeta)$ - Default choice $d = \{0, 1\}$ - First update of beliefs $\tilde{\zeta}(d, \zeta)$	- Temporary exclusion from debt markets - Output cost $\phi_j(y)$	- Trading in SM A - Choice of b' and $\tilde{\pi}$ - Message $m = \{L, NL\}$ - Second update of beliefs $\hat{\zeta}(m, \tilde{\zeta})$ - Trading in SM B	- Primary markets open: Coupon payments & debt issuance b'

window around the government's report of inflation. Our model, however, is calibrated at quarterly frequency. Because our goal is to use this elasticity to discipline the learning parameter α, in order to address this frequency disconnect, we extend our baseline model by allowing different trading instances within the same period.

Figure A.2 describes the timing assumption under this extension. There are two instances of trading in secondary markets (SM) within a period. The first instance (A) is at the beginning of stage 1, right after the government's default decision and before message m is realized. The second (B) occurs after lenders observe message m and update their beliefs. In both cases, SM bond prices are cum dividend and thus include the expected dividend payments at the end of period t, when the primary market (PM) opens.

Let $q_A(y, b, \tilde{\zeta})$ denote the pricing kernel at trading instance A. This price depends on the expected value of the bond at trading instance B, once message m is realized but before coupons are paid. It is given by

$$q^{(A)}\left(y, b, \tilde{\zeta}\right) = \tilde{\zeta} q_C^{(A)}\left(y, b, \tilde{\zeta}\right) + \left(1 - \tilde{\zeta}\right) q_S^{(A)}\left(y, b, \tilde{\zeta}\right), \tag{A.10}$$

where, for each $j = \{C, S\}$,

$$q_j^{(A)}\left(y, b, \tilde{\zeta}\right) = \sum_{M=\{L,NL\}} Prob\left(m = M \mid \tilde{\Pi}_j^\star\right) q^{(B)}\left(y, b, \hat{\zeta}(m)\right), \tag{A.11}$$

where $\tilde{\Pi}_j^\star \equiv \tilde{\Pi}_j^\star(y, b, \tilde{\zeta})$, $\hat{\zeta}(m)$ is given by Equation (4) and $q_B(y, b, \hat{\zeta}(m))$ is the price of a bond at trading instance B. This price, in turn, is given by

$$q^{(B)}\left(y, b, \hat{\zeta}(m)\right) = \left\{\lambda + (1 - \lambda)\left(z + q(y, b'^\star, \zeta'\right)\right\}, \tag{A.12}$$

where $b^{\star\prime} \equiv b^{\star\prime}(y, b, \tilde{\zeta})$ is the bond policy, $\zeta' \equiv \zeta'\left(\hat{\zeta}(m)\right)$ is given by Equation (5), and $q(\cdot)$ is the price of a bond in the primary market. For any b', it is given by

$$q(y, b', \zeta') = \frac{1}{1+r} \int_y \left\{ \zeta' \left(1 - d_C^{\star\prime}\right) q_C^{(A)}\left(y', b', \tilde{\zeta}'\right) + (1 - \zeta')\left(1 - d_S^{\star\prime}\right) q_S^{(A)}\left(y', b', \tilde{\zeta}'\right) \right\} dF(y' \mid y), \tag{A.13}$$

where $d_j^{\star\prime} \equiv d_j^{\star}(y', b', \zeta')$ refers to the conjectured next-period default choice for type j and the posterior $\tilde{\zeta}'$ is given by $\tilde{\zeta}' \equiv \tilde{\zeta}(d' = 0, \zeta'; d_S^{\star\prime}, d_C^{\star\prime})$ (as defined in Equations (3)). Most importantly, notice that by replacing Equations (A.10)-(A.12) in Equation (A.13), we obtain the pricing equation of the baseline model (defined in Equation (A.8)). In other words, the proposed extension nests our baseline model.

This model extension allows us to compute the intraperiod change in bond prices before and after the realization of message m. Given our timing assumption regarding the choice of the bond policy, the realized message m does not affect debt issuances. Thus, changes in bond prices between trading instances A and B are purely driven by changes in a government's reputation.

Lastly, we introduce a measure of the BE inflation rate in the model. To this end, we first compute the price of an auxiliary inflation-indexed bond (IIB) with the same maturity structure as b, but whose payoffs depend on the government's misreport. The pricing kernel of this IIB is analogous to that of the nominal bond, with the only difference being that the (expected) bond payments are adjusted by $(1 + \tilde{\Pi}_j^{\star})$. The model-implied BE inflation rates for trading instances (A) and (B) are defined as

$$BE^{(A)}\left(y, b, \tilde{\zeta}\right) = \text{Yield}^{(A)}\left(y, b, \tilde{\zeta}\right) - \text{Yield}_{\text{IIB}}^{(A)}\left(y, b, \tilde{\zeta}\right)$$
$$BE^{(B)}\left(y, b, \hat{\zeta}(m)\right) = \text{Yield}^{(B)}\left(y, b, \hat{\zeta}(m)\right) - \text{Yield}_{\text{IIB}}^{(B)}\left(y, b, \hat{\zeta}(m)\right),$$

where the bond yields can be computed directly from the pricing kernels. We can then compute intraperiod price changes between instances A and B as follows:

$$\Delta BE(m) = BE^{(B)}\left(y, b, \hat{\zeta}(m)\right) - BE^{(A)}\left(y, b, \tilde{\zeta}\right) \tag{A.14}$$

$$\Delta \ln SP(m) = \ln SP^{(B)}\left(y, b, \hat{\zeta}(m)\right) - \ln SP^{(A)}\left(y, b, \tilde{\zeta}\right). \tag{A.15}$$

Conditional on the state of the economy, these model-implied changes only depend on the realization of message m. In this regard, they resemble the high-frequency measures for ΔBE and $\Delta\ln SP$ that we compute in our empirical analysis. A final issue to consider is that the realized m, in turn, depends on the optimal choice of $\tilde{\pi}$ (which is an endogenous object). That is, both $\Delta BE(m)$ and $\Delta \ln SP(m)$ are endogenous variables. In the data, our estimation approach was precisely chosen to address this reverse-causality concern. In the model, we can isolate the causal effect of the misreport on spreads by constructing a counterfactual in which we shock the optimal misreport policy by $\epsilon_{\tilde{\pi}}$. This shock affects the realization of message m and hence the posterior $\hat{\zeta}(m)$ and prices. Let m be the realized message under the optimal $\tilde{\pi}^{\star} \equiv \tilde{\pi}_S\left(y, b, \tilde{\zeta}\right)$ policy and let m_ϵ be the realized message under a counterfactual in which the misreport is $\tilde{\pi}^{\star} + \epsilon_{\tilde{\pi}}$. Our model-implied elasticity is then defined as

$$\eta_{BE,SP} \equiv \mathbb{E}\left[\frac{\Delta \ln SP\left(m_\epsilon\right) - \Delta \ln SP\left(m\right)}{\Delta BE\left(m_\epsilon\right) - \Delta BE\left(m\right)}\right]. \tag{A.16}$$

In the quantitative analysis, we calibrate the learning parameter α so that our model-implied elasticity, $\eta_{BE,SP}$, matches the one in our empirical analysis.

Appendix B. Empirical Analysis

B.1. *Data Sources*

Data on Argentina's official inflation rate are obtained from the *Instituto Nacional de Estadística y Censos* (INDEC). Actual report dates are obtained from historical articles posted online by the newspaper *La Nación*. Data on Argentine consumption and debt are obtained from national sources. Data on bond yields and bond characteristics are obtained from Bloomberg. Data on Argentina's stock index (MERVAL) and forward contracts for the Argentine peso are also obtained from Bloomberg. For global control variables (used throughout the paper), we retrieve the VIX index, the S&P 500 index, and the MSCI Emerging Markets ETF index from Datastream.

B.2. *List of Event Days*

Table B.1 lists all of the days on which the Argentine government reported the inflation rate between 2007 and 2010. To construct the list, we accessed historical articles from the Argentine newspaper *La Nación* by using the tool provided by the Wayback Machine.[42] The table also displays the announced (monthly) inflation rate.

Table B.1. Reporting Dates

Event	Month	Reported Day	Rate (%)	Event	Month	Reported Day	Rate (%)
1	Jan-07	2/5/2007	1.14	25	Jan-09	2/11/2009	0.53
2	Feb-07	3/5/2007	0.30	26	Feb-09	3/11/2009	0.43
3	Mar-07	4/11/2007	0.77	27	Mar-09	4/14/2009	0.64
4	Apr-07	5/4/2007	0.74	28	Apr-09	5/13/2009	0.33
5	May-07	6/5/2007	0.42	29	May-09	6/11/2009	0.33
6	Jun-07	7/5/2007	0.44	30	Jun-09	7/14/2009	0.42
7	Jul-07	8/7/2007	0.50	31	Jul-09	8/12/2009	0.62
8	Aug-07	9/7/2007	0.59	32	Aug-09	9/4/2009	0.83
9	Sep-07	10/5/2007	0.80	33	Sep-09	10/14/2009	0.74
10	Oct-07	11/6/2007	0.68	34	Oct-09	11/12/2009	0.80
11	Nov-07	12/6/2007	0.85	35	Nov-09	12/11/2009	0.83
12	Dec-07	1/7/2008	0.93	36	Dec-09	1/15/2010	0.93
13	Jan-08	2/7/2008	0.93	37	Jan-10	2/12/2010	1.04
14	Feb-08	3/6/2008	0.47	38	Feb-10	3/12/2010	1.25
15	Mar-08	4/10/2008	1.13	39	Mar-10	4/14/2010	1.14
16	Apr-08	5/9/2008	0.83	40	Apr-10	5/12/2010	0.83
17	May-08	6/10/2008	0.56	41	May-10	6/14/2010	0.75
18	Jun-08	7/11/2008	0.64	42	Jun-10	7/14/2010	0.73
19	Jul-08	8/11/2008	0.37	43	Jul-10	8/13/2010	0.80
20	Aug-08	9/11/2008	0.47	44	Aug-10	9/15/2010	0.74
21	Sep-08	10/10/2008	0.51	45	Sep-10	10/15/2010	0.72
22	Oct-08	11/11/2008	0.43	46	Oct-10	11/12/2010	0.84
23	Nov-08	12/10/2008	0.34	47	Nov-10	12/16/2010	0.73
24	Dec-08	1/13/2009	0.34	48	Dec-10	1/14/2011	0.84

[42]See https://archive.org/web/.

B.3. *Argentina's Fundamentals*

During the period of study, Argentina's fundamentals were in line with those of the region. The left panel of Figure B.1 shows that Argentina's GDP growth showed a behavior similar to that observed in other Latin American countries. If anything, Argentina was growing faster than the rest of the region before the GFC. The right panel of Figure B.1 shows that the dynamics of the stock market were also aligned with those of the region. Although not shown, Argentina's stock of debt was on a downward trend since 2006.

Figure B.1. Argentina versus LATAM countries

(A) GDP growth (real, year-year in %) (B) Stocks growth (USD, year-year in %)

Notes: The figure shows the real GDP growth (Panel A) and the average stock-market return (Panel B) of Argentina and other Latin American countries (Brazil, Chile, Colombia, Mexico, Peru, and Uruguay). Thicker lines show the Argentine case. The gray areas show the GFC period.

B.4. *Analysis of Bond Yields and Break-even Inflation Rate*

We provide additional details on the Argentine government's bond yields and on the construction of the BE inflation rate. Table B.2 shows static information for the Argentine bonds, whose daily data was retrieved from Bloomberg for the period 2007-2012. The top panel describes the nominal bonds in the dataset (both dollars and pesos) and the bottom panel reports the inflation-indexed bonds (IIBs).

Table B.2. Static Information for Argentina's Bonds

(a) Dollar-denominated Bonds

ISIN	Maturity	Currency	Coupon Frequency
ARARGE03F482	12jun2012	ARS	S/A
ARARGE03F243	28mar2011	USD	S/A
ARARGE03F342(*)	12sep2013	USD	S/A
ARARGE03F144(*)	03oct2015	USD	S/A
ARARGE03F441	17apr2017	USD	S/A
US040114GL81	31dec2033	USD	S/A
US040114GK09	31dec2038	USD	S/A

(b) Inflation-linked Bonds

ISIN	Maturity	Currency	Coupon Frequency
ARARGE03B309	15mar2014	ARS	Monthly
ARARGE03E931(*)	30sep2014	ARS	S/A
ARARGE035162	03jan2016	ARS	Monthly
ARBNAC030255	04feb2018	ARS	Monthly

Notes: The table shows static information for all of the bonds in our sample. The top panel shows information for nominal bonds (both dollars and pesos). The bottom panel shows information for IIBs. Bonds with an asterisk (*) are the ones used in the main analysis.

We use the yields of these bonds to compute a measure of the BE inflation. Let $\text{Yield}^{\$}_{m,t}$ be the annualized yield of a nominal bond (in pesos) with maturity m. Let $\text{Yield}^{\text{IIB}}_{m,t}$ be the yield of an IIB with maturity m. The BE inflation rate is defined as

$$BE_{m,t} = \text{Yield}^{\$}_{m,t} - \text{Yield}^{\text{IIB}}_{m,t}.$$

A major setback is that only three nominal bonds denominated in pesos were actively trading during the period considered. Moreover, there is only one bond for which we have yields data during 2007, and the first observation is in July (i.e., 6 months after the government started misreporting the inflation rate). To circumvent this issue, we construct a measure for the BE rate using the yields of nominal bonds in dollars ($\text{Yield}_{m,t}^{US\$}$) and the expected devaluation of the peso, as implied by forward currency contracts. Let F_0 denote the spot exchange rate. Let F_m be the future exchange rate m months from today. Let $\delta_m^e \equiv \frac{F_m - F_0}{F_0}$ be the expected devaluation rate in m-periods. We can then compute the BE inflation rate as

$$BE_{m,t} = \text{Yield}_{m,t}^{US\$} - \text{Yield}_{m,t}^{\text{IIB}} + \delta_m^e. \tag{B.1}$$

Ideally, to compute the BE rate we need to consider bonds with the same maturity and frequency of coupon payments. From Table B.2, notice that all of the nominal bonds pay coupons on a semi-annual frequency. Only one IIB pays coupons at this frequency (highlighted with an asterisk). This is the bond we use in our main analysis. To compute the BE rate, we then use the average yield for the two dollar-denominated bonds whose maturities are closest to this IIB.[43]

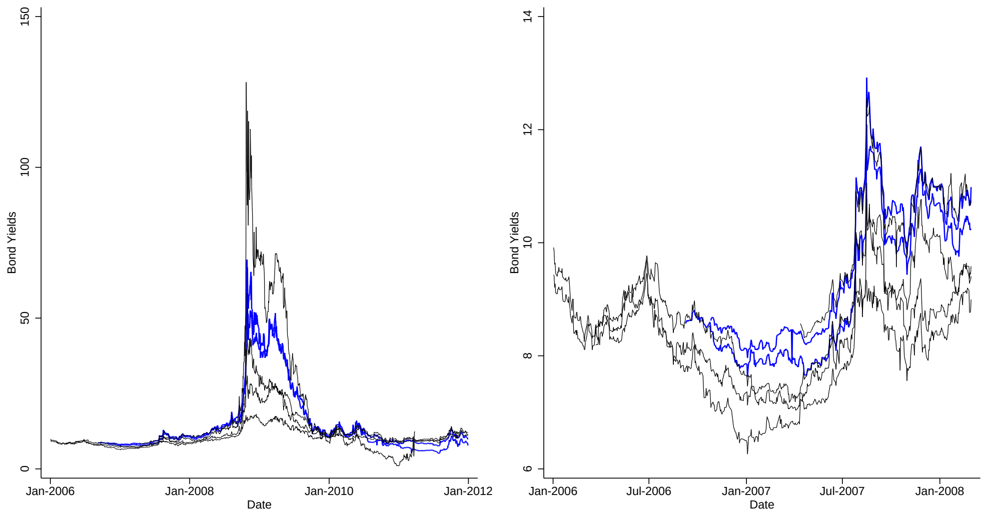
The top panel of Figure B.2 shows annual yields for the dollar-denominated bonds.[44] The bottom panel shows yields for the IIBs. Blue lines depict the bonds used in our main analysis. The left panels show yields for the period 2006-2012 and the right panels focus on the pre-GFC period. Overall, all of the different yields move in tandem, particularly in the pre-GFC period. Figure B.3 shows different measures for the BE inflation rate. In all of the cases depicted, we use the IIB with semi-annual payments. Thus, each line of Figure B.3 corresponds to a different dollar-denominated bond. The blue line shows the measure of the BE inflation rate used in our main analysis. Overall, all of the measures strongly co-move during the sample period.

[43] Results are robust to using different dollar-denominated bonds.

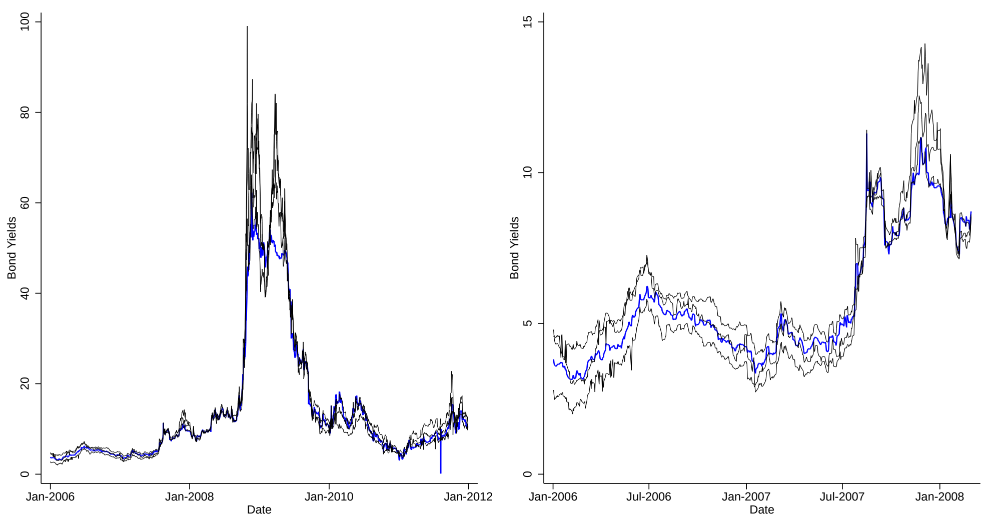
[44] Yields for the last two dollar-denominated bonds in Table B.2 are omitted because the maturities of these bonds are significantly larger.

Figure B.2. Yield of Argentina's Bonds

(a) Dollar-denominated Bonds

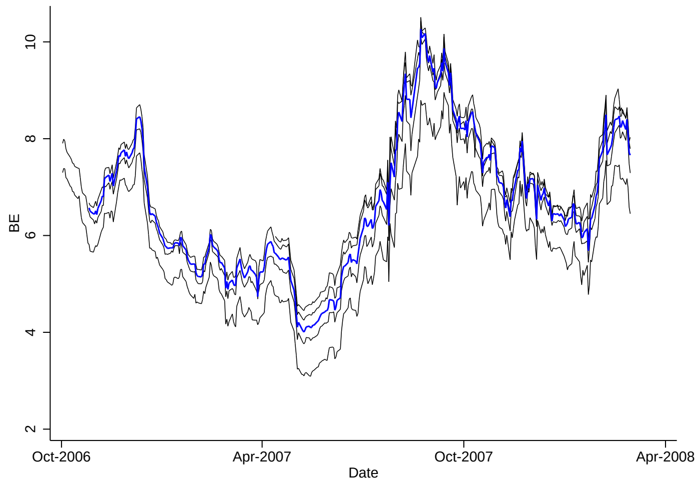


(b) Inflation-indexed Bonds



Notes: The figure shows the annual yields for different dollar-denominated bonds and inflation-linked bonds issued by Argentina's national government. The blue line corresponds to the bonds used in the main analysis. Left panels include the period 2006-2012. Right panels zoom in on the pre-GFC period.

FIGURE B.3. Break-even Inflation Rates



Notes: The figure shows different measures of the break-even inflation rate. The blue line corresponds to the measure used in the main analysis.

B.5. *Inflation Misreport and Changes in the BE inflation rate*

Throughout our analysis, we use changes in the BE inflation rate around inflation announcements as a high-frequency proxy for the unexpected component of the misreport. To assess whether this is a reasonable assumption, we compare the daily change in the BE rate (around the day on which the government reported the inflation) with the monthly change in the *observed* misreport. The underlying idea is that, to the extent that agents learn gradually from previous misreports, the change in the observed misreport should be informative about the unexpected component. We define the observed change as

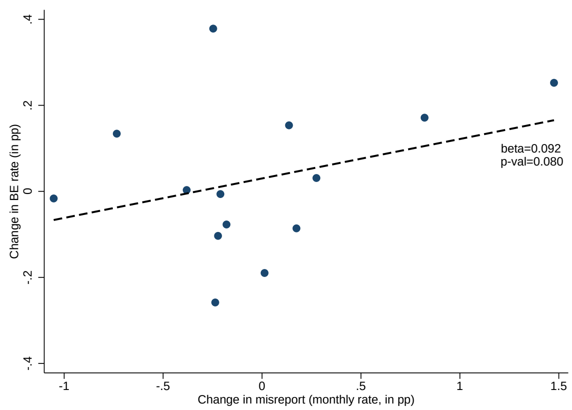
$$\epsilon_{\tilde{\pi},t} \equiv \tilde{\pi}^o_t - \tilde{\pi}^o_{t-1}, \tag{B.2}$$

where $\tilde{\pi}^o_t$ is the observed inflation misreport at time t —i.e., the difference between the inflation announced by the government and the one obtained from alternative (private) sources.

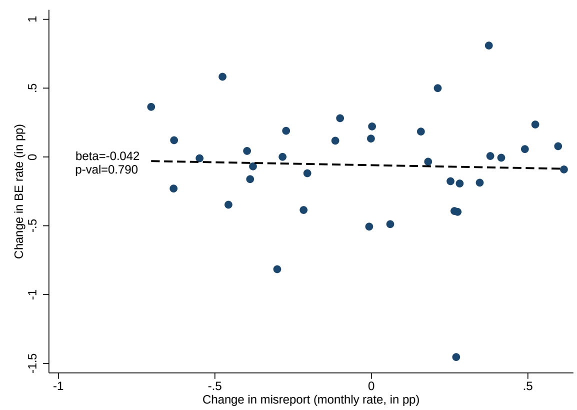
Figure B.4 compares the change in the BE rate with the change in the observed misreport. Panel (A) shows a positive (and significant) relation between these two variables prior to March 2008. However, Panel (B) shows that this positive relation disappears post March 2008, which suggests that agents were no longer surprised and the misreport was already priced. Although the analysis is qualitative in nature, we take this as evidence to suggest that the change in the BE rate is a good proxy for the unexpected component of the inflation misreport.

FIGURE B.4. Changes in Break-even Inflation Rate and Inflation Misreport

(A) Pre March 2008



(B) Post March 2008



Notes: The figure shows the change in the BE inflation rate versus the change in the *observed* inflation misreport. We define the observed misreport as the difference between the inflation announced by the government and that provided by alternative (i.e., private) sources. Panel (A) corresponds to the sample period prior to March 2008 and Panel (B) to the period post March 2008.

B.6. *Discussion: Exchange Rate Risk*

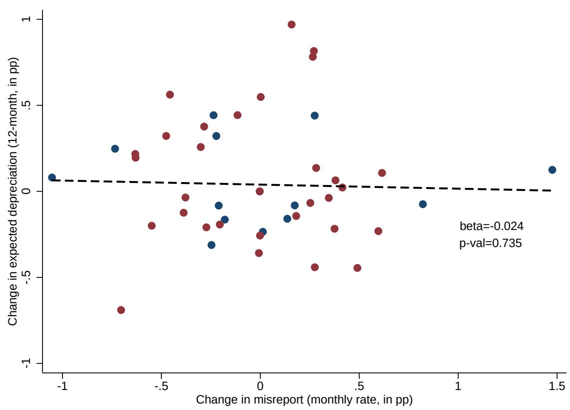
A concern of using the expected depreciation rate to construct the BE inflation rate (as shown in Equation B.1) is the presence of exchange-rate risk. That is, the dynamics of the BE rate around days on which the government announces the inflation could be driven mostly by adjustments in the expected depreciation rate.

To address this concern, Figure B.5 compares the change in the expected (12-month) depreciation rate with the monthly change in the inflation misreport ($\epsilon_{\tilde{\pi},t}$, our proxy for the unexpected component of the misreport). Panel (A) shows the results when using currency forward contracts to compute the daily change in the expected depreciation rate around days in which the government reported the inflation rate. The results suggest that there is no relation between these two measures.

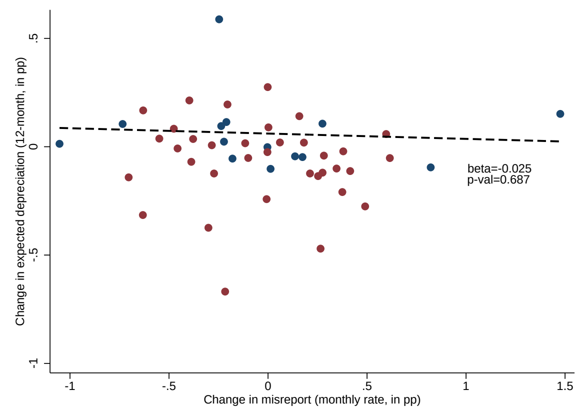
An additional issue to consider is that currency forward contracts tend to be quite volatile. This implies that the volatility of the expected depreciation rate computed from these forwards can be relatively high, and this could be driving the low correlation with changes in the observed misreport. To tackle this point, in Panel (B) we consider a 5-day moving average for the forward-implied expected depreciation rate. Even after this smoothing, there is still no relation between

FIGURE B.5. Changes in Expected Depreciation Rate and Inflation Misreport

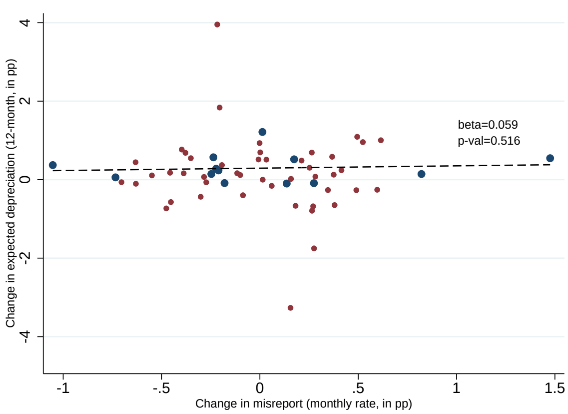
(A) Forwards



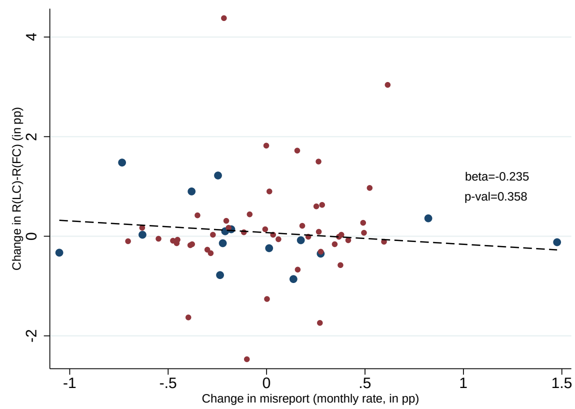
(B) Forwards (5-day Moving Average)



(C) Expectations Survey



(D) Deposit Rate Differentials



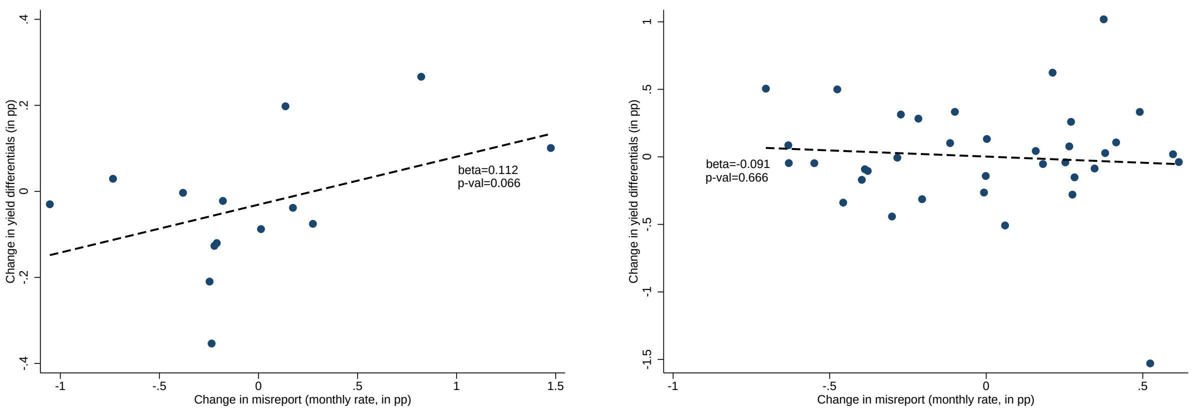
Notes: The figure shows the change in the expected 12-month depreciation rate versus the change in inflation misreport. Panels (A) and (B) show the change in the expected depreciation based on currency forward contracts. Panel (C) uses a measure of expected depreciation based on a weekly survey conducted by the Argentine Central Bank. Panel (D) computes a monthly measure of expected depreciation based on the difference between local- and foreign-currency time deposit rates. In all of the cases, blue (red) dots correspond to the sample period prior (post) March 2008. Best fit line is for the sample period prior to March 2008.

these two variables. To avoid the complications that may arise due to the volatility of currency forward contracts, we use this moving average in our baseline analysis.

We complement the previous analysis with two lower-frequency measures of the expected depreciation rate. First, we use data on nominal exchange rate expectations based on a survey conducted by the Argentine Central Bank at weekly frequency. Professional forecasters (e.g. banks, hedge funds, brokers, think tanks, and universities) are asked about their expected

Figure B.6. Changes in Yields Differential and Inflation Misreport

(A) Pre March 2008 (B) Post March 2008



Notes: The figure shows the change in the yield differential between dollar-denominated and IIB bonds versus the change in the *observed* inflation misreport. We define the observed misreport as the difference between the inflation announced by the government and that provided by alternative (i.e., private) sources. Panel (A) corresponds to the sample period prior to March 2008. Panel (B) shows the post March 2008 period.

exchange rate for the current and following calendar years. We use the cross-sectional average response to construct a weekly time series of the expected 12-month depreciation rate. Panel (C) of Figure B.5 compares this measure with the monthly change in the observed misreport. Results suggest there is no relation between these two variables.[45] In Panel (D), we show the expected depreciation computed from the difference between local- and foreign-currency time deposit rates at monthly frequency—obtained from the Argentine Central Bank. Results are similar to the previous cases.

Finally, Figure B.6 shows the relation between the change in the observed misreport and the change in the yields differential, $\text{Yield}^{US\$}_{m,t} - \text{Yield}^{\text{IIB}}_{m,t}$. For the pre March 2008 sample, there is a positive relation between these two variables. The relation vanishes for the later part of our sample period.

To sum up, the results from this appendix suggest that out of the two components behind our measure of the BE rate (as defined in Equation (B.1)), the yields differential is the one that co-moves with the inflation misreport.

[45] It is not necessarily the case that survey responses occur after the government's announcement of inflation. In 73% of the cases, however, the surveys were conducted the week after the inflation report. Results are similar if only those months are included in the figure.

B.7. *Test of Identifying Assumption*

We present an F-test to verify the main assumption of the Rigobon and Sack approach—namely, that the variance of the shocks to ΔBE_t is larger on event days. As it can be seen from Equation (14) in the main text, the Rigobon and Sack instrument is relevant only under the assumption that $\Lambda \equiv \sigma_{\eta,E}/\sigma_{\eta,NE} > 1$. To test this, we conduct a hypothesis test in which $\sigma(\Delta BE)_E = \sigma(\Delta BE)_N$. Our one-sided alternative hypothesis is that $\sigma(\Delta BE)_E > \sigma(\Delta BE)_N$. The F-tests reported in Table B.3 strongly reject the hypothesis of equal variances, which provides evidence in favor of $\Lambda > 1$. A bias-corrected stratified bootstrap shows that we can also reject the hypothesis of equal variances. Although not reported, the tests are not significant during and after the GFC.

Table B.3. Test of Identifying Assumption

	Window 1	Window 2	Window 3	Window 4
Window Type				
Event	2-day window	3-day window	2-day window	3-day window
Non-event	All other days	All other days	4-day window	4-day window
Standard Deviation				
Event	0.294	0.265	0.294	0.265
Non-event	0.189	0.178	0.165	0.165
Ratio Test: $\sigma_{\Delta BE,E} > \sigma_{\Delta BE,NE}$				
F-test				
F-value	2.432	2.214	3.197	2.602
$P(F > f)$	0.000	0.000	0.000	0.001
BC Bootstrap - One-Sided CI				
90% CI Lower Bound	1.222	1.198	1.352	1.244
95% CI Lower Bound	1.116	1.120	1.221	1.163

Notes: The top panel reports the standard deviations of the daily change in the BE inflation rate across different event and non-event windows. The bottom panel shows two tests for the equality of variances of changes in the BE rate. We include the results for a traditional F-test and a bias-corrected bootstrap. Sample period: January 2007-February 2008.

B.8. *Robustness Analysis*

In this section, we present a robustness analysis for our main empirical exercise of Section 3. We consider OLS regressions and a standard event-study analysis based on narrow windows around the inflation announcement.

The analysis in this section relies on a stronger identifying assumption compared to the heteroskedasticity-based identification analysis in the main text. In particular, it requires that changes in Argentina's BE inflation rate during event windows are driven exclusively by the inflation announcement. Both the OLS and event-study estimates are thus subject to the concern that other (potentially unobserved) common factors may have changed during those event days. Another problem is the smaller sample size, since we only focus on days around the inflation announcement. Nevertheless, the analysis is still useful to further study the relation between ΔBE_t and $\Delta \ln SP_t$ around days on which the government announces the inflation rate, and to analyze how this relation changes across time.

OLS Estimates

We start by describing the regression behind Figure 5. We consider the following specification:

$$\Delta \ln SP_t = \alpha_0 + \alpha_1^E \Delta BE_t \times \mathbb{I}_{t,E} + \alpha_1^{NE} \Delta BE_t \times \mathbb{I}_{t,NE} + \alpha_2 F_t + \epsilon_t,$$

where $\mathbb{I}_{t,E}$ and $\mathbb{I}_{t,NE}$ are indicators for event and non-event days (based on a 2-day window around the inflation announcement), and F_t is a vector of global controls (as described in the main text). To construct the figure, we run the previous specification across different days of our sample based on a 12-month rolling window.

For the rest of this subsection, we split our sample between event and non-event days, and for each set of events, we consider the following specification:

$$\Delta \ln SP_t = \alpha_0 + \alpha_1 \Delta BE_t + \alpha_2 F_t + \epsilon_t.$$

Panel (A) of Table B.4 shows OLS estimates for our baseline sample period (January 2007-February 2008). When we focus on narrow windows around the announcement of inflation, the estimates are negative and significant (and in line with those presented in the main text). However, for non-event days, the OLS estimates are not significant. This suggests that outside of announcement days, changes in sovereign spreads are unrelated to changes in the BE rate. Although the Argentine government kept misreporting the inflation rate after 2008, the results are not significant once we exclude the first year of the sequence of misreports (Panel B).[46] Through the lens of our reputational model, these results suggest that the lenders' prior ζ reached its lower bound after the first year of the misreports. In other words, the market was no longer surprised.

[46]Although not reported, the results are not significant either for the period 2008-2009. This may not be surprising, given that changes in Argentina's sovereign spreads during the GFC may have mostly been driven by external factors.

Table B.4. OLS Estimates

(A) January 2007-February 2008

	(1)	(2)	(3)	(4)	(5)
Event Window	Full Sample	2-day Window		3-day Window	
ΔBE	-1.726*	-6.691***	-0.473	-6.783***	-0.666
Standard Error	(0.972)	(1.998)	(0.981)	(1.787)	(1.040)
Observations	258	24	234	36	219
Days Included	All	Event Days	Non-Event Days	Event Days	Non-Event Days
Controls	Yes	Yes	Yes	Yes	Yes

(B) January 2010-February 2011

	(1)	(2)	(3)	(4)	(5)
Event Window	Full Sample	2-day Window		3-day Window	
ΔBE	-0.184	-0.485	-0.132	-0.719	-0.034
Standard Error	(0.278)	(0.894)	(0.298)	(0.632)	(0.325)
Observations	259	25	234	39	219
Days Included	All	Event Days	Non-Event Days	Event Days	Non-Event Days
Controls	Yes	Yes	Yes	Yes	Yes

Notes: The table shows results for the OLS estimators. The dependent variable is $\Delta \ln SP_t$. Panel (A) shows estimates for the period January 2007-February 2008. Panel (B) shows results for the period January 2010-February 2011. The first column includes all of the days in the sample. The other columns only include 2- and 3-day windows around the inflation announcement. Controls include the VIX index, the S&P 500 index, and the MSCI Emerging Markets ETF index. Robust standard errors are reported in parentheses. ***, **, and * denote significance at 1%, 5%, and 10%, respectively.

Event-study Results

We next present a standard event-study analysis to estimate the effect of the misreports on Argentina's sovereign spreads. Let NE denote the set of non-event days and $L = \mid NE \mid$. We first estimate a factor model for the non-event-days,

$$\Delta \ln SP_t = \phi_0 + \phi_1 F_t + \nu_t,$$

where F_t is the same vector of global controls used in the main analysis. We then use those estimates to generate a time series of abnormal changes in Argentina's sovereign spreads and

Table B.5. Event-study Approach

Event Type	# Events	Obs	$\Delta ln(\bar{S}P^A)$	J1-stat	$\Delta\bar{B}E$
		2007-2008			
Good News Event	6	12	-2.229	-2.935	0.187
Bad News Event	5	10	1.477	1.775	-0.105
		2010-2011			
Good News Event	5	10	-0.251	-0.400	0.283
Bad News Event	7	14	0.761	1.439	-0.208

Notes: The table shows the results for the event-study analysis. Events are classified as good or bad news based on the average change in the BE rate around the Argentine government's report of inflation. The top panel shows results for January 2007-February 2008. The bottom panel shows results for January 2010-February 2011. $\Delta\ln\bar{S}P^A$ denotes the average across $\sum_{t\in k}\Delta\ln SP_t^A$.

estimate its variance (assuming that errors are homoskedastic). That is,

$$\Delta\ln SP_t^A = \Delta\ln SP_t - \hat{\phi}_0 - \hat{\phi}_1 F_t$$
$$\hat{\sigma}^2_{SP} = \frac{1}{L}\sum_{t\in NE}(\Delta\ln SP_t^A)^2.$$

Next, we classify our event windows into two categories depending on the observed change in the BE inflation rate. Let $\mu^{E,j}_{\Delta BE}$ be the average ΔBE_t for event window j, and $\mu^{NE}_{\Delta BE}$ be the average for non-event days. From the pool of event days we create two categories:[47]

(1) If $\mu^{E,j}_{\Delta BE} < \mu^{NE}_{\Delta BE}$, we label the event window j as a bad news event (BNE).
(2) If $\mu^{E,j}_{\Delta BE} > \mu^{NE}_{\Delta BE}$, we label the event window j as a good news event (GNE).

In the first category, for instance, the drop in the BE inflation rate during event window j is larger than the average change for non-event days. This can be interpreted as an increase in the unexpected underreport of inflation, and thus a bad news event.

For each category $k = \{BNE, GNE\}$, we compute the cumulative abnormal change across all of the events of the same type k: $CA(SP)_k = \sum_{t\in k}\Delta\ln SP_t^A$. Notice that CA(SP) adds abnormal changes across different windows (i.e., non-consecutive days). Finally, we report the

[47] Ideally, we would like to have three categories: bad news, no news, and good news. Given our small sample, we decided to focus only on two broad categories. Results are similar if we classify events based on the median change (instead of on the mean change).

following $J1$ statistic described in Campbell et al. (1997):

$$J1_k = \frac{CA(SP)_k}{\sqrt{L_k \times \hat{\sigma}^2_{SP}}},$$

where $L_k = \mid E_k \mid$ denotes the total number of days for each type of event k. Under the null hypothesis that misreport events have no effect on $\Delta \ln SP$, $J1_k$ is asymptotically distributed as a standard normal variable. The problem is that there are few observations in each category and, therefore, asymptotic normality may be a poor approximation. Thus, the results should be interpreted only as suggestive evidence.

Table B.5 reports the results based on a 2-day window. For the period January 2007-February 2008 (top panel), there is an asymmetric effect of changes in the BE rate on spreads. The average (daily) change in (log) spreads is 1.5% and −2.2% in the bad and good news event, respectively. For the period January 2010-February 2011, the effects are smaller in magnitude and not significant.[48] The results are consistent with our reputational channel and in line with those presented in the main text.

B.9. *A Reputational Channel*

We provide further evidence that supports the reputation channel. As a starting point, we extend our baseline model and allow for the possibility that the inflation misreport can directly affect the real economy (Equations (17)-(19) in the main text). For convenience, we replicate that system of equations below:

$$\Delta BE_t = \beta_0 + \beta_1 \Delta \ln SP_t + \beta_2 R_t + \beta_3 X_t + \eta_t \tag{B.3}$$

$$\Delta \ln SP_t = \alpha_0 + \alpha_1 \Delta BE_t + \alpha_2 R_t + \alpha_3 X_t + \epsilon_t \tag{B.4}$$

$$R_t = \gamma_0 + \gamma_1 \Delta BE_t + \gamma_3 X_t + \nu_t, \tag{B.5}$$

where we assume that η_t, ϵ_t, ν_t, and X_t are uncorrelated. Substituting Equation (B.5) into Equations (B.3) and (B.4), it is straightforward to show that

$$\Delta BE_t \left(1 - \beta_2\gamma_1\right) = \left(\beta_0 + \beta_2\gamma_0\right) + \beta_1 \Delta \ln SP_t + \left(\beta_3 + \beta_2\gamma_3\right) X_t + \left(\eta_t + \beta_2\nu_t\right)$$

$$\Delta \ln SP_t = \left(\alpha_0 + \alpha_2\gamma_0\right) + \left(\alpha_1 + \alpha_2\gamma_1\right) \Delta BE_t + \left(\alpha_3 + \alpha_2\gamma_3\right) X_t + \left(\epsilon_t + \alpha_2\nu_t\right).$$

Under the same set of assumptions as in Section 3.3, and under the additional assumption that $\sigma_{\nu,E} = \sigma_{\nu,NE}$, although we cannot identify α_1, our identification strategy allows us to identify $\tilde{\alpha_1} \equiv \alpha_1 + \alpha_2\gamma_1$. The coefficient α_1 would account for our "reputational channel," while $\alpha_2\gamma_1$ accounts for a "fundamentals channel" —i.e., the effect of inflation announcements on sovereign

[48] Although not reported, the effects for 2008-2009 are not significant either.

spreads through the real economy. Thus, to the extent that $\alpha_2 \neq 0$ and $\gamma_1 \neq 0$, our baseline estimates for α_1 would be biased. In what follows, we analyze the sign and magnitude of the bias.

According to the sovereign debt literature, we would expect α_2 to be negative: A fall in economic activity (as proxied by stock-market returns) should increase a country's default risk. The sign of γ_1 is a priori unclear because there may be different channels through which the inflation announcements end up affecting the economy's fundamentals. First, the misreports may lead to distortions in relative prices, increase uncertainty, and reduce investment, which may end up decreasing Argentina's economic activity. Second, changes in the BE rate may be capturing not only news regarding the misreports but also news about the "true" inflation rate. For a high-inflation country such as Argentina, $\Delta BE_t < 0$ may thus be perceived as a good signal about the fundamentals of the economy, which should have a positive effect on R_t.

Although we cannot identify α_1, under the system of equations (B.3)-(B.5), we can identify the γ_1 parameter. To see this, substitute Equation (B.4) into (B.3) to get the following system of equations:

$$\Delta BE_t \left(1 - \beta_1 \alpha_1\right) = \left(\beta_0 + \beta_1 \alpha_0\right) + \left(\beta_2 + \beta_1 \alpha_2\right) R_t + \left(\beta_3 + \beta_1 \alpha_3\right) X_t + \left(\eta_t + \beta_1 \epsilon_t\right)$$

$$R_t = \gamma_0 + \gamma_1 \Delta BE_t + \gamma_3 X_t + \nu_t.$$

From this expression, it is clear that our set of identifying assumptions allows us to identify γ_1. Table 3 (in the main text) shows our estimates. Across all of the specifications, the point estimates for γ_1 are not statistically significant. Therefore, the inflation announcements do not seem to have a direct effect on the Argentine stock market. We take this as further evidence that supports our reputational channel.

We end our discussion of possible biases by considering the case in which ΔSP_t could affect R_t—as Hébert and Schreger (2017) find. To do this, we consider the following system of equations:

$$\Delta BE_t = \beta_0 + \beta_1 \Delta \ln SP_t + \beta_2 R_t + \beta_3 X_t + \eta_t \tag{B.6}$$

$$\Delta \ln SP_t = \alpha_0 + \alpha_1 \Delta BE_t + \alpha_2 R_t + \alpha_3 X_t + \epsilon_t \tag{B.7}$$

$$R_t = \gamma_0 + \gamma_1 \Delta BE_t + \gamma_2 \Delta \ln SP_t + \gamma_3 X_t + \nu_t, \tag{B.8}$$

where we have replaced Equation (B.5) with (B.8). According to Hébert and Schreger (2017), we should expect a negative effect of sovereign spreads on stock returns (i.e., $\gamma_2 < 0$).

Under our identifying assumptions, although we cannot identify α_1 or γ_1, it is easy to show that we can identify $\tilde{\alpha_1} \equiv \frac{\alpha_1+\alpha_2\gamma_1}{1-\alpha_2\gamma_2}$ and $\tilde{\gamma_1} \equiv \frac{\gamma_1+\alpha_1\gamma_2}{1-\alpha_2\gamma_2}$. The biases are given by $Bias(\alpha_1) = \alpha_2\tilde{\gamma}_1$ and $Bias(\gamma_1) = \gamma_2\tilde{\alpha_1}$, respectively. Based on our reputational mechanism and the findings in Hébert and Schreger (2017) (i.e., $\gamma_2 < 0$), we argue that the bias for γ_1 should be positive. Moreover, the fact that our point estimates for $\tilde{\gamma_1}$ (i.e., those reported in Table 3) are close to zero and not statistically significant suggests that the bias for α_1 is small. Since our estimate for γ_1 is positively biased, we argue that our estimates for α_1 are also likely to be positively biased (given $\alpha_2 < 0$). Hence, in terms of magnitudes, we could interpret them as a lower bound. In Appendix B.10, we analyze these points more formally through the lens of a structural VAR.

B.10. *Identified Structural VAR*

We estimate a structural VAR that incorporates the interactions between the inflation misreports, spreads, and economic activity. We identify structural shocks to the misreport policy using high-frequency changes in the BE inflation rate, and we study their effects on sovereign spreads and the real economy.

Let $\boldsymbol{Y}_t \equiv \left(Y_t^p, \boldsymbol{Y}_t^{q\prime}\right)'$, where Y_t^p is the policy variable and $\boldsymbol{Y}_t^q$ denotes the rest of the variables of the VAR. Consider the following structural and reduced form VAR:

$$\text{Structural Form:} \quad \boldsymbol{A}\boldsymbol{Y}_t = \sum_{j=1}^{p} \boldsymbol{C}_j \boldsymbol{Y}_{t-j} + \boldsymbol{\epsilon}_t \tag{B.9}$$

$$\text{Reduced Form:} \quad \boldsymbol{Y}_t = \sum_{j=1}^{p} \boldsymbol{B}_j \boldsymbol{Y}_{t-j} + \boldsymbol{u}_t, \tag{B.10}$$

where $\boldsymbol{u}_t = \boldsymbol{S}\boldsymbol{\epsilon}_t$, $\boldsymbol{S} = \boldsymbol{A}^{-1}$, and $\boldsymbol{B}_j = \boldsymbol{A}^{-1}\boldsymbol{C}_j$. The vectors $\boldsymbol{\epsilon}_t$ and $\boldsymbol{u}_t$ represent structural and reduced-form shocks, respectively. Let ϵ_t^p be the structural policy shock. In our analysis, this is an inflation misreport shock. Let $\boldsymbol{s}$ denote the column in $\boldsymbol{S}$ associated with ϵ_t^p. The response of the endogenous variables to a shock to the inflation-misreport is given by

$$\boldsymbol{Y}_t = \sum_{j=1}^{p} \boldsymbol{B}_j \boldsymbol{Y}_{t-j} + \boldsymbol{s}\epsilon_t^p.$$

This means that, given estimates for $\{\boldsymbol{B}_j\}_{j=1}^p$, we only need to identify $\boldsymbol{s}$ to compute the impulse response. To this end, we follow an instrumental approach similar to that of Mertens and Ravn (2013) and Gertler and Karadi (2015).

The method consists of finding a vector of instruments $\boldsymbol{Z}_t$ such that $\mathbb{E}\left[\boldsymbol{Z}_t\epsilon_t^p\right] \neq \boldsymbol{0}$ and $\mathbb{E}\left[\boldsymbol{Z}_t\boldsymbol{\epsilon}_t^{q\prime}\right] = \boldsymbol{0}$, where $\boldsymbol{\epsilon}_t^q$ is the vector of structural shocks other than the policy shock ϵ_t^p. Given a vector of instruments, the procedure for obtaining estimates of $\boldsymbol{s}$ can be decomposed in two steps. First, we obtain estimates of $\boldsymbol{B}_j$ and $\boldsymbol{u}_t$ from an OLS regression of the reduced form VAR. Second, we identify $\boldsymbol{s}$ using the estimated reduced-form residuals and the vector of instruments $\boldsymbol{Z_t}$. Let u_t^p be the reduced form residuals associated with the policy equation, and let $\boldsymbol{u}_t^q$ be the reduced form residuals from the other equations. Let $\boldsymbol{s}^q \in \boldsymbol{s}$ be the response of $\boldsymbol{u}_t^q$ to a unit increase in ϵ_t^p. As explained in Gertler and Karadi (2015), we can obtain an estimate of $\boldsymbol{s}^q$ and s^p from a two-stage OLS estimation. In the first stage, we regress u_t^p onto $\boldsymbol{Z}_t$ to get $\hat{u}_t^p$. In the second stage, we regress $\boldsymbol{u}_t^q$ onto $\hat{u}_t^p$ to obtain estimates for $\boldsymbol{s}^q$ and s^p.

We consider the policy variable Y_t^p to be the inflation misreport. An additional complication in our application is that Y_t^p is not perfectly observable, since market participants do not observe the true inflation rate. We assume that agents observe $\tilde{\boldsymbol{Y}}_t = \boldsymbol{Y}_t + \boldsymbol{\eta}_t$, where $\boldsymbol{\eta}_t = (\eta_t, \mathbf{0}')'$. We assume that η_t is i.i.d. and orthogonal to $\boldsymbol{Y}_\tau$ for any τ. Being measurement errors, we also assume that $\mathbb{E}[\boldsymbol{\eta}_t \epsilon_t^p] = 0$, $\mathbb{E}\left[\boldsymbol{\eta}_t \boldsymbol{\epsilon}_t^{q\prime}\right] = \mathbf{0}$, and $\mathbb{E}[\boldsymbol{\eta}_t \boldsymbol{Z}_t'] = \mathbf{0}$. Although strong, these are sufficient conditions to identify our parameters of interest. Our reduced form VAR can be written as $\tilde{\boldsymbol{Y}}_t = \sum_{j=1}^p \tilde{\boldsymbol{B}}_j \tilde{\boldsymbol{Y}}_{t-j} + \tilde{\boldsymbol{u}}_t$. This is the same expression as the one in Equation (B.10) with $\boldsymbol{u}_t = \tilde{\boldsymbol{u}}_t - \boldsymbol{\eta}_t + \sum_{j=1}^p \tilde{\boldsymbol{B}}_j \boldsymbol{\eta}_{t-j}$ and $\boldsymbol{B_j} = \tilde{\boldsymbol{B}}_{\boldsymbol{j}}$. Since $\boldsymbol{Y}_{t-j} \perp \boldsymbol{\eta}_{t-\tau}$, the OLS estimator would actually return an unbiased estimate for $\boldsymbol{B}_j$. The structural form VAR is $\tilde{\boldsymbol{A}}\tilde{\boldsymbol{Y}}_t = \sum_{j=1}^p \tilde{\boldsymbol{C}}_j \tilde{\boldsymbol{Y}}_{t-j} + \tilde{\boldsymbol{\epsilon}}_t$, which is the same as Equation (B.9) with $\boldsymbol{\epsilon}_t = \tilde{\boldsymbol{\epsilon}}_t - \tilde{\boldsymbol{A}}\boldsymbol{\eta}_t + \sum_{j=1}^p \tilde{\boldsymbol{C}}_j \boldsymbol{\eta}_{t-j}$, $\boldsymbol{A} = \tilde{\boldsymbol{A}}$, and $\boldsymbol{C}_j = \tilde{\boldsymbol{C}}_j$. Given the orthogonality assumptions on $\boldsymbol{\eta}_t$, we have

$$\mathbb{E}[\boldsymbol{Z}_t \tilde{\epsilon}_t^p] = \mathbb{E}\left[\boldsymbol{Z}_t \left(\epsilon_t^p + \tilde{\boldsymbol{A}}_{[1,1]}\eta_t - \sum_{j=1}^p \tilde{\boldsymbol{C}}_{[1,1],j}\eta_{t-j}\right)\right] = \mathbb{E}[\boldsymbol{Z}_t \epsilon_t^p].$$

Thus, we can still use $\boldsymbol{Z}_t$ to identify the structural shock to the misreport equation.

We estimate the previous SVAR using monthly data. Let $\tilde{\boldsymbol{Y}}_t \equiv (M_t, SP_t, IP_t)'$, where M_t is the noisily observed misreport, SP_t are sovereign spreads, and IP_t is an indicator of Argentina's economic activity. We define M_t as the difference between the alternative measures of the inflation rate and the inflation rate reported by the Argentine government—as shown in Figure 3. For Argentina's sovereign spreads, we take the residual of a projection of daily spreads (in logs) onto the set of factors used in Section 3.3 (VIX, SP, and EEM) to control for global conditions.[49] We then compute the median value for each month. Our measure of economic activity is the "Estimador Mensual de Actividad Economica" as reported by the INDEC. This is a seasonally adjusted monthly variable that captures Argentine nonfinancial economic activity. We take the residual of the projection of this index onto the following set of external variables:

[49]We do not introduce these global variables into the VAR because it would significantly increase the number of coefficients to estimate and yield a relatively small number of observations.

oil price, US unemployment rate, and the US 10-year Treasury yield. Lastly, we consider changes in the BE rate around inflation announcements, ΔBE_t, to be our instrument for the identification of the structural misreport policy shocks.[50]

For the first step of the procedure, we use monthly data for the period 2006-2010 to estimate the reduced form VAR. Given the small number of observations, we only choose one lag for the VAR. For the second step, we use data on ΔBE_t for the period between February 2007 and September 2008 to identify the vector $\boldsymbol{s}$.[51] It could be the case that the previous sufficient conditions for η_t may not hold. For instance, to the extent that the consumption baskets considered in the official and alternative measures of inflation differ, the dynamics for the observed misreports may have an important seasonal component. To control for this, we seasonally adjust the observed misreports before introducing them into our VAR specification. It could also be the case that the volatility of the observed misreports depends on its level. To mitigate this concern, we normalize the misreports by the official level of inflation.

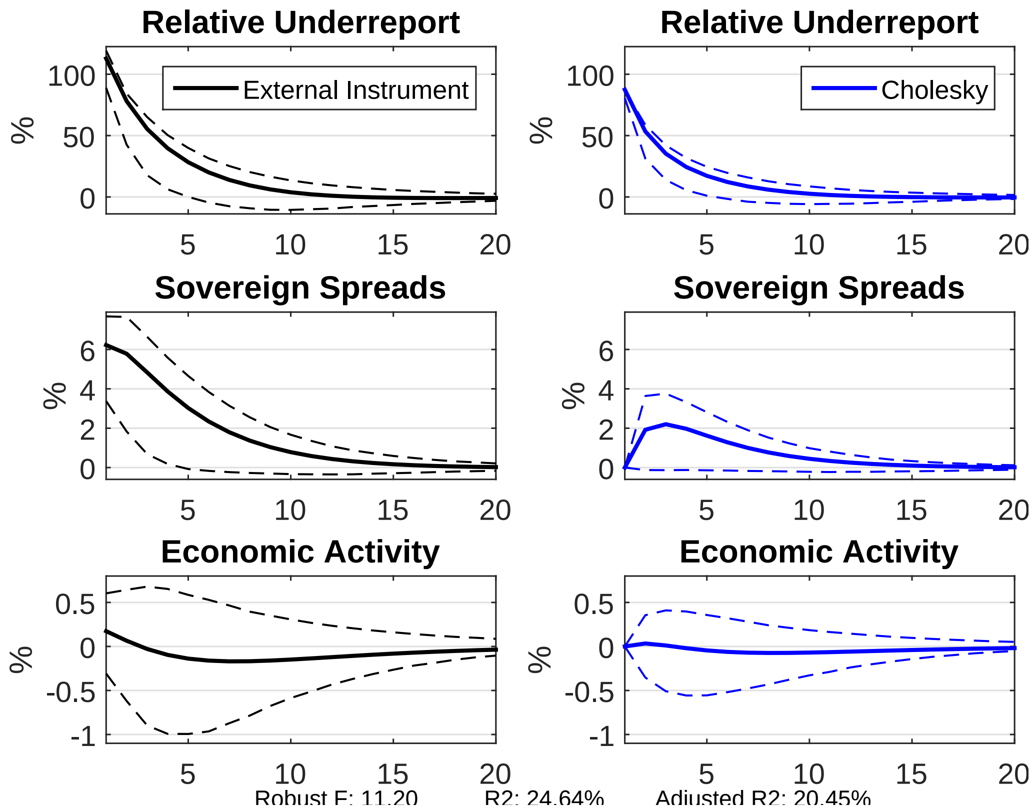
Figure B.7 shows the results. The left panel shows the response of inflation underreport, spreads, and economic activity upon a 1–sd structural shock to the misreport policy. We find that an unexpected underreport leads to an increase in spreads, but it does not have a sizable effect on economic activity. The robust F-test statistic is greater than 10, which suggests that the external instrument is valid (see Stock et al., 2002). For comparison, the right panel shows the responses based on a Cholesky decomposition.[52] The responses are similar to the ones of our identified SVAR—albeit smaller in magnitude. Overall, the results in Figure B.7 are in line with those presented in the main text. Despite some shortcomings (small sample size, lower frequency), we take this analysis as further evidence to support our reputational mechanism.

[50]We compute ΔBE_t in a short window around the inflation announcement, which mitigates the concern that the instrument might be correlated with other structural shocks.

[51]We include a larger sample period than in our main analysis of Section 3 to increase the number of monthly observations. Our results are qualitatively similar if we use data until the GFC. However, those results are less precisely estimated due to a reduction in the number of observations. Similarly, we include data for ΔBE_t until September 2008 (instead of March 2008, as in our main analysis) to avoid having a small number of monthly observations and a weak instrument problem. Results are similar if we consider $\Delta \ln BE_t$ as our instrument.

[52]The assumed (decreasing) order of exogeneity is: (i) economic activity, (ii) spreads, and (iii) misreports.

Figure B.7. Impulse Response to a Misreport Shock



Notes: This figure shows the response of inflation underreport, spreads, and economic activity to a 1–sd structural shock to misreport. Dashed lines denote the 90% confidence interval based on a Wild Bootstrap.

Appendix C. Quantitative Analysis

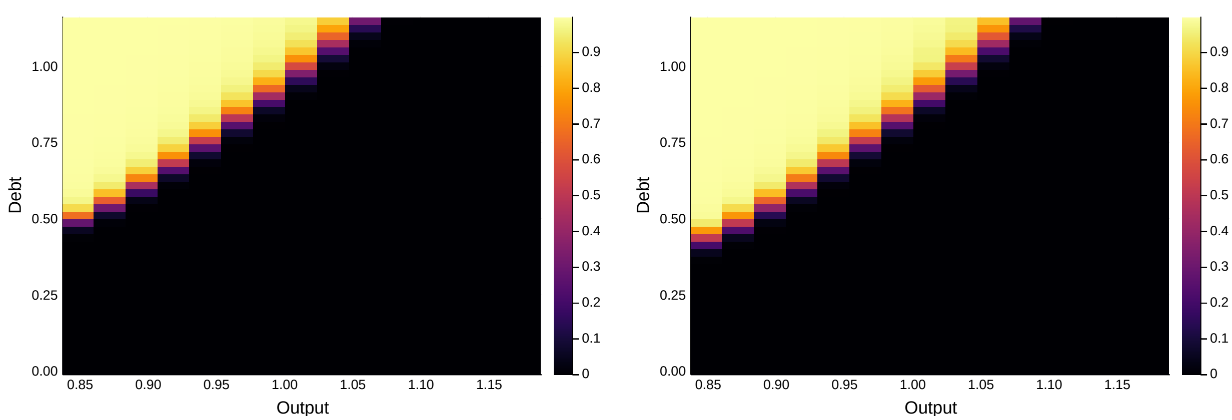
This appendix complements our main quantitative analysis. First, we present figures that describe the optimal default policy and the bond-pricing kernel. Second, we describe how the expected posterior and our model-implied elasticity vary with different values for the learning parameter α. Third, we show that our main results hold for different persistence values of government types. Fourth, we compare our baseline model with a perfect information counterfactual and provide a welfare analysis. Lastly, we provide additional figures for the Argentine counterfactual and describe the model's solution method.

C.1. *Default Policies and Bond Prices*

Figure C.1 shows the optimal default-repayment policy for the C-type (Panel A) and for the S-type (Panel B). The figure assumes a relatively high reputation ($\zeta = 0.8$). The area in the upper-left corner represents points of the state space in which the j-type defaults on its debt ($d = 1$). These are states in which debt is high and output is low. Notice that, by assumption, the default set for the S-type is slightly larger than that for the C-type.

Figure C.1. Default - Repayment Sets

(A) Default Set - C-type (B) Default Set - S-type

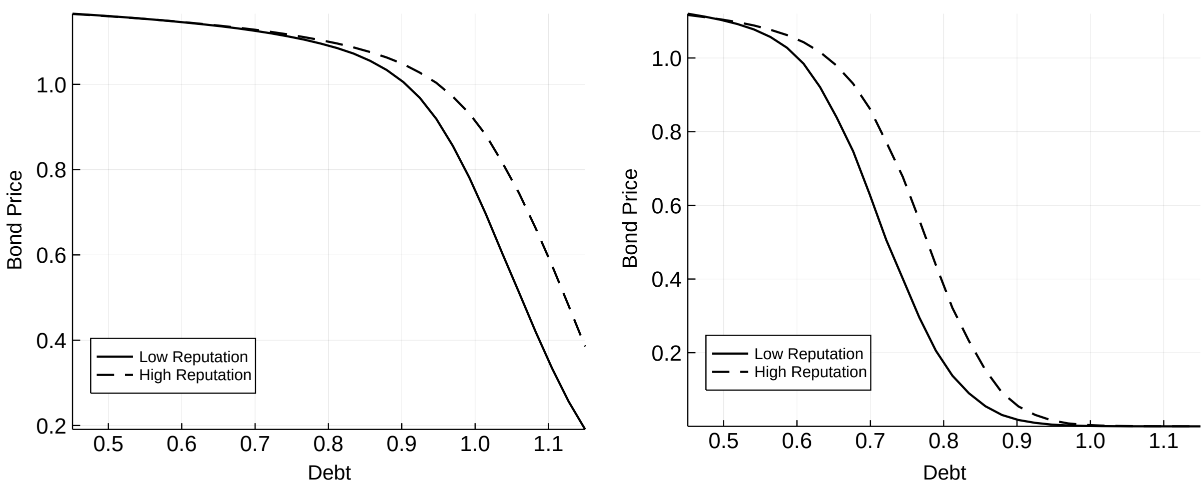


Notes: The figure shows the default/repayment sets for the C-type (left panel) and for the S-type (right panel). The area in the upper-left corner of each panel represents the points of the state space in which the government defaults ($d = 1$).

The different default sets imply that changes in ζ' affect the lenders' perceived probability of a default, and thus the government's borrowing costs. The effect is larger as we move closer to the default boundary. This is illustrated in Figure C.2. The figure shows the bond-pricing kernel $q(y, b', \zeta')$ for different values of b', ζ', and y. The effect of reputation on bond prices

Figure C.2. Bond-pricing Kernel

(A) High Output (B) Low Output



Notes: The figure shows the bond-pricing kernel $q(y, b', \zeta')$ for different combinations of b' and ζ'. The left (right) panel shows the case when output is high (low).

clearly depends on the macro fundamentals (b', y). For instance, for $b' = 0.7$, changes in ζ' have almost no effect on the bond price if output is high. When output is low, however, as the government is closer to its default boundary, changes in ζ' can have a sizable effect on bond prices.

C.2. *Analysis of the Learning Parameter α*

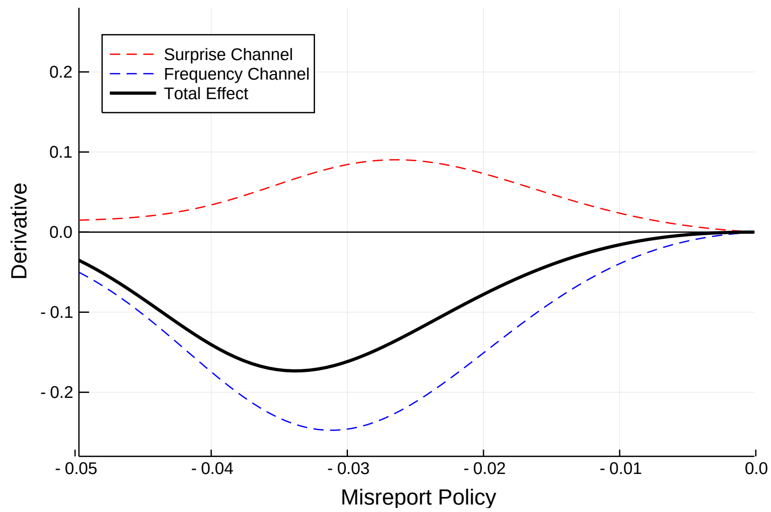
We analyze how changes in the learning parameter α affect the expected posterior reputation and the model-implied elasticity $\eta_{BE,SP}$. We start by analyzing the effects of α on the expected posterior. At the beginning of stage 1, before message m is realized and for a given choice of $\tilde{\pi}$, the expected posterior can be written as

$$\mathbb{E}_m\left(\hat{\zeta}(m)\right) = \Gamma\left(\tilde{\pi}; \sigma, \alpha\right)\hat{\zeta}_L + \left(1 - \Gamma\left(\tilde{\pi}; \sigma, \alpha\right)\right)\hat{\zeta}_{NL},$$

where, with a slight abuse of notation, $\hat{\zeta}_m \equiv \hat{\zeta}\left(m, \tilde{\zeta}; \tilde{\Pi}^\star_S, \sigma, \alpha\right)$ for $m = \{L, NL\}$ denotes the updated posterior conditional on the realized message m (as defined in Equation (4)). Taking derivatives with respect to α around $\tilde{\pi} = \tilde{\Pi}^\star_S$, we get

$$\frac{\partial \mathbb{E}_m\left(\hat{\zeta}(m)\right)}{\partial \alpha}\Bigg|_{\tilde{\pi}=\tilde{\Pi}^\star_S} = \underbrace{\Gamma'_\alpha\left(\tilde{\Pi}^\star_S; \sigma, \alpha\right)\left(\hat{\zeta}_L - \hat{\zeta}_{NL}\right)}_{\text{Frequency Channel}} + \underbrace{\Gamma\left(\tilde{\Pi}^\star_S; \sigma, \alpha\right)\frac{\partial \hat{\zeta}_L}{\partial \alpha} + \left[1 - \Gamma\left(\tilde{\Pi}^\star_S; \sigma, \alpha\right)\right]\frac{\partial \hat{\zeta}_{NL}}{\partial \alpha}}_{\text{Surprise Channel}}. \tag{C.1}$$

FIGURE C.3. The Frequency and Surprise Channels



Notes: The figure shows a decomposition of $\frac{\partial E(\hat{\zeta}(m))}{\partial \alpha}$ into a frequency channel (blue dashed lines) and a surprise channel (red dashed lines). The black solid line shows the total effect. We assume a prior of $\tilde{\zeta} = 0.5$. Derivatives are evaluated at $\tilde{\pi} = \tilde{\Pi}_S^\star$. We divide by 100 to interpret the results as the effect of a 1 pp change in α.

We label the first term as a *frequency channel*. By design, a larger α (i.e., a lower $\mid \alpha \mid$) weakly increases the probability that $m = L$ is realized (i.e., $\Gamma'_\alpha(\cdot) \geq 0$). Since $\hat{\zeta}_L \leq \hat{\zeta}_{NL}$, this channel implies that a larger α weakly decreases the expected posterior.

We define the last term as a *surprise channel*, which has two components. First, a lower α implies that for any $\tilde{\Pi}_S^\star < 0$, it is less likely to get $m = L$; but upon observing such message, it is very informative about the government's being of the S-type. Second, a larger α implies that for any $\tilde{\Pi}_S^\star < 0$, it is less likely to get $m = NL$; but upon receiving such message, it is more informative about the government's being of the C-type. For the case in which $\Gamma(.)$ is given by the CDF of a normal distribution, it is easy to show that $\frac{\partial \hat{\zeta}_L}{\partial \alpha}$ and $\frac{\partial \hat{\zeta}_{NL}}{\partial \alpha}$ are positive for any $\tilde{\Pi}_S^\star < 0$.

From the previous analysis, it is then clear that the frequency channel has an opposite effect to the surprise channel. In Figure C.3, we provide a numerical example that illustrates the magnitude of each channel. For all $\tilde{\pi} \in [\underline{\pi}, 0]$, the magnitude of the frequency channel is always larger, and thus the total effect is negative.[53]

[53]For values of $\tilde{\pi} < -0.06$, the frequency channel is negligible (since the probability of receiving message $m = L$ is almost one) and the total effect is positive—albeit small. Such values, however, are outside our grid for $\tilde{\pi}$, since they imply an annualized underreport of inflation higher than 24%, which is even larger than the Argentine inflation rate.

We now turn to the effects of α on the elasticity $\eta_{BE,SP}$ (as defined in Equation (20)). Suppose that we shock the optimal misreport policy by $\epsilon_{\tilde{\pi}}$. Let m be the realized message under the optimal $\tilde{\pi}^{\star} \equiv \tilde{\pi}^{\star}\left(y, b, \tilde{\zeta}\right)$, and let m_{ϵ} be the realized message in a counterfactual in which the misreport is $\tilde{\pi}^{\star} + \epsilon_{\tilde{\pi}}$. Based on a first-order Taylor expansion of $\eta_{BE,SP}$ for a small $\epsilon_{\tilde{\pi}}$ shock around $\tilde{\pi}^{\star} = \tilde{\Pi}^{\star}_{S}$, and after taking derivatives with respect to α, we get

$$\frac{\partial \eta_{BE,SP}}{\partial \alpha}\frac{1}{\mid \eta_{BE,SP} \mid} = \underbrace{(-)\frac{\partial \phi\left(\frac{\alpha-\tilde{\Pi}^{\star}_{S}}{\sigma}\right)}{\partial \alpha}\frac{1}{\phi\left(\frac{\alpha-\tilde{\Pi}^{\star}_{S}}{\sigma}\right)}}_{\text{Frequency Channel}} + \underbrace{\frac{\partial \Delta\left(SP/BE\right)}{\partial \alpha}\frac{1}{\mid \Delta\left(SP/BE\right) \mid}}_{\text{Surprise Channel}}, \tag{C.2}$$

where $\Delta\left(SP/BE\right) \equiv \frac{\ln SP(m=L)-\ln SP(m=NL)}{BE(m=L)-BE(m=NL)} < 0$ and $\phi\left(\cdot\right)$ denotes the density function of a standard normal variable.

The frequency channel captures how α affects $\eta_{BE,SP}$ through its effects on the change in the probability of receiving message $m = L$ given the $\epsilon_{\tilde{\pi}}$ shock. It is easy to show that the frequency channel is given by $\frac{\alpha-\tilde{\Pi}^{\star}_{S}}{\sigma^2}$. Hence, it is negative as long as $\alpha <\tilde{\Pi}^{\star}_{S}$. Given our calibration of the model, this is almost always the case in our simulations. The surprise channel captures the effect of α on the way in which changes in m affect spreads and the BE rate. Although we cannot characterize this channel (since both SP and BE are endogenous objects), based on a numerical derivative analysis, we find that it is typically negative and much smaller (in magnitude) than the frequency channel.[54] Overall, we find that a larger α leads to a lower $\eta_{BE,SP} < 0$.

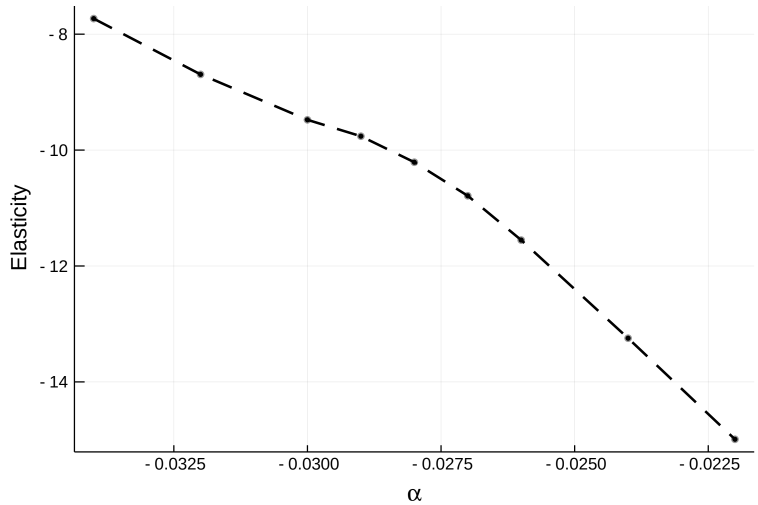
The previous analysis applies for just a small perturbation around $\tilde{\Pi}^{\star}_{S}$. Thus, it may fail to capture the nonlinearities of our model. To account for the potential implications of these nonlinearities, we solve the model for different values of α while keeping all other model parameters fixed. Figure C.4 shows the results. In line with the previous analysis, we find that the magnitude of $\eta_{BE,SP}$ monotonically increases with α. We argue that this monotonicity allows us to identify α and discipline our model.

C.3. *The Persistence of Government Types*

In our main analysis, we set the transition matrix across types, T, to reflect an election cycle of 8 years. Identifying the persistence of a government type is challenging, since it is a low frequency parameter. In fact, other papers in the literature target a persistence between 2 and 16 years (see, for example, D'Erasmo, 2011; Amador and Phelan, 2021; and Fourakis, 2021). Our baseline calibration is somewhere in between that range.

[54] In our simulations, the frequency channel accounts for more than three quarters of the total effect.

FIGURE C.4. The Relation between α and $\eta_{BE,SP}$



Notes: The figure shows the model-implied elasticity, $\eta_{BE,SP}$, for different values of the learning parameter α.

TABLE C.1. Reputation Premium and Persistence of Government Types

Moment	Description	Persistence		
		Low	Baseline	High
$\mathbb{E}[\Upsilon]$	Average reputation premium	64bp	98bp	101bp
$\mathbb{E}[\Upsilon/SP]$	Incidence reputation premium on spreads	8%	13%	14%
$\sigma(\Upsilon)/\sigma(SP)$	Reputation premium volatility	35%	44%	43%
$\sigma(SP\|\zeta_H)/\sigma(SP)$	Spread volatility under high reputation	68%	60%	61%
$\mathbb{E}[\Upsilon/SP\|Y < Y_l]$	Incidence with low output	16%	21%	22%
$\text{corr}(\Upsilon, \log Y)$	Correlation reputation premium & output	−62%	−64%	−62%
$\text{corr}(\Upsilon/SP, \log Y)$	Correlation reputation incidence & output	−67%	−67%	−64%

Notes: The table shows moments related to the reputation premium, Υ, and the link between Υ and the economy's fundamentals. Each column reports the results for different persistence values (i.e., Markov transition T).

In Table C.1, we analyze the implications of alternative persistence values on the reputation premium. We consider an average change across types every 4 and 12 years. For each case, we recalibrate the parameters $\{\beta, \bar{\chi}_0, \bar{\chi}_1, \bar{\chi}_2\, B, \alpha\}$ to match the same set of target moments as those of Table 5. The table shows that our results are robust to different parameterizations. In particular, the reputation premium accounts for a large share of spreads, especially if output is low.

C.4. *Comparison with the Perfect-information Case*

We further compare the implications of our baseline model with respect to a case in which the type of government is perfectly observable. We first analyze the implications in terms of spreads and borrowing costs. We then analyze the welfare costs of information frictions.

Figure C.5 shows the ratio of spreads between the baseline model and the perfect-information case for different values of debt (b, x-axis), output (y, line shifts), and reputation (ζ, shaded areas). We focus on periods in which the government is of the $C-$type. The bounds of the areas are based on the interquartile range of simulated values of ζ conditioning on periods in which the government is of the $C-$type. The figure shows that once we fix a value of debt, spreads under the baseline model can be significantly larger than under the perfect-information counterfactual, particularly if output is low. For instance, when output is just half a standard deviation below its mean and for the average debt of the C-type (denoted with a vertical dotted line), spreads under the imperfect-information case can be up to 30% higher.

For our welfare analysis, we compute how much we would need to compensate the government (in consumption units) for it to be indifferent between the baseline model and the perfect-information case. Since preferences are CRRA, the certainty equivalent consumption (CEC) is given by $\omega(y,b,z) = \left[\frac{W^{PI}(y,b)}{W(y,b,z)}\right]^{\frac{1}{1-\gamma}} - 1$, where $W(\cdot)$ denotes the value function under the baseline model, and $W^{PI}(\cdot)$ is the value function under the alternative scenario. A positive value of $\omega(y,b,z)$ means that the government is better off in the perfect-information case.[55]

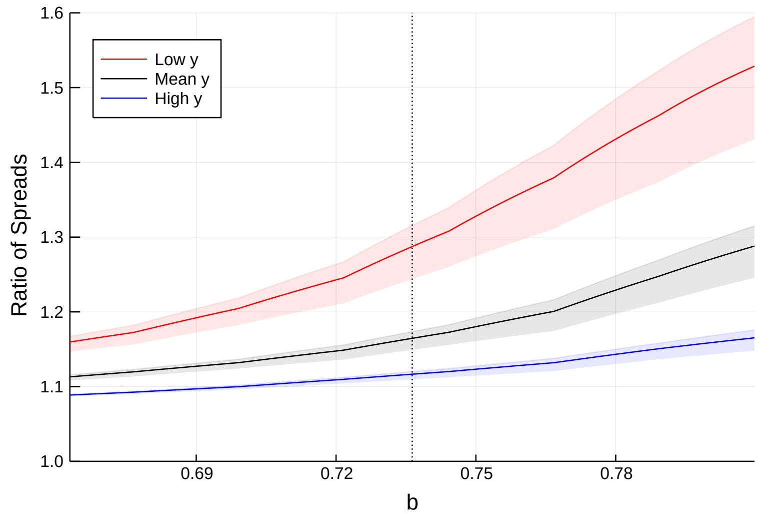
Figure C.6 shows the results for different combinations of (b,ζ). The CEC measure is always positive. More importantly, the CEC is larger in points of the state space in which reputation is low and debt is high. This is because in those points of the state space, the C-type faces significantly larger borrowing costs, since it is not able to perfectly reveal its type.

Lastly, we analyze the welfare implications associated with the presence of alternating types. Figure C.7 shows the CEC that makes the j-type indifferent between the baseline model (with imperfect information and alternating types) and a case in which types are fixed (and observable).[56] Panel (A) shows that the C-type is significantly worse off in the baseline scenario. Panel (B), on the other hand, shows that the S-type is better off in the baseline model (the CEC is negative). This is because under imperfect information, the S-type can attain a larger level of debt and lower borrowing costs.

[55] We define W and W^{PI} as the average between the value functions for the C- and S-type.

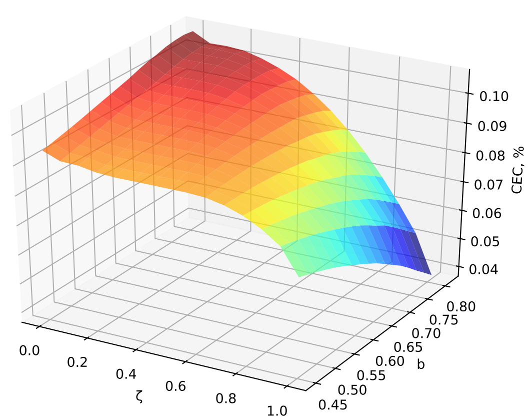
[56] The analysis has the caveat that preferences are different under the fixed-type counterfactual, since each j-type is not subject to changes in preferences over default.

FIGURE C.5. Ratio of Spreads: Baseline Model versus Perfect-information Case



Notes: The figure shows the ratio of spreads between the baseline model with imperfect information and the perfect-information case. Each line displays the ratio of spreads for a particular level of output, y. Low (high) y corresponds to half a standard deviation below (above) the average y. The shadowed areas display the ratio of spreads across different values of ζ. We include a wide range of ζ that covers the 25th-75th percentile of the simulated path of ζ once we condition on periods in which the government is of the C-type. The vertical line denotes the average debt in our baseline model.

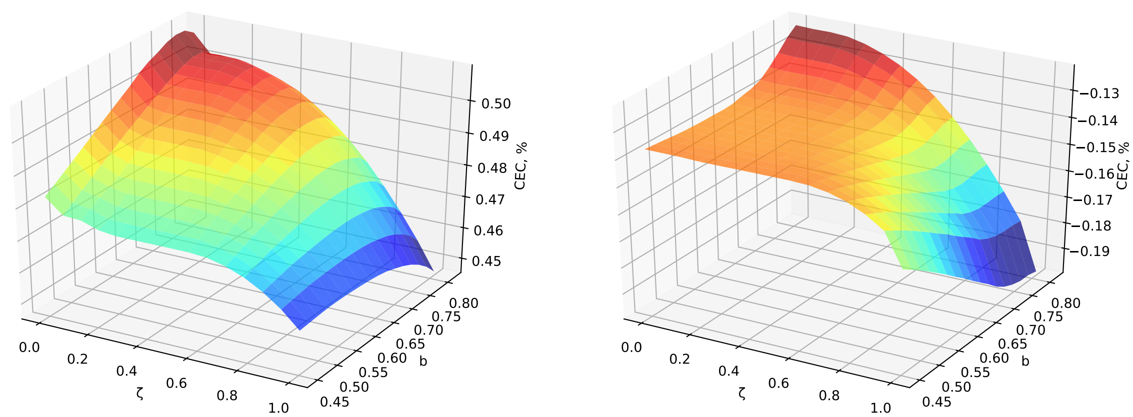
FIGURE C.6. CEC - Baseline Model vs. Perfect-information Case



Notes: The figure shows the additional certainty equivalent consumption (CEC) that makes the government indifferent between the baseline model (with imperfect information) and a case in which the type of government is perfectly observable. The figure assumes that output is at its mean.

FIGURE C.7. CEC - Baseline Model vs. Fixed-types Case

(A) C-type (B) S-type



Notes: The figure shows the additional CEC that makes the j-type indifferent between the baseline model (with imperfect information) and a case in which government types are fixed (and observable). Panel (A) shows the results for the C-type. Panel (B) shows the results for the S-type. The figure assumes that output is at its mean.

C.5. *The Argentine Case: Additional Material*

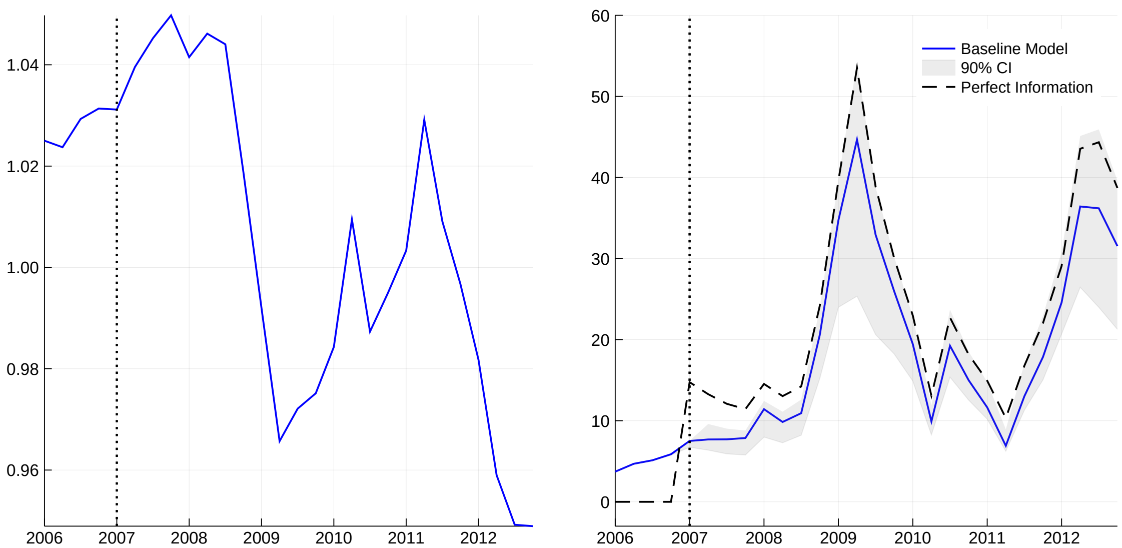
In this section, we provide additional figures for the Argentine simulations of Section 4.4. Figure C.8a shows the path for output used for the simulations. Results are based on a log-linear trend. Argentina underwent a severe financial crisis during 2001-2002 that produces a structural break in the log-linear trend. Therefore, we estimate the log-linear trend with a structural break starting in 2002. Results are almost identical if we instead consider the HP cycle of output. We opted for the log-linear cycle to keep consistency with the model calibration and computation of moments in Section 4.1.

Figure C.8b compares the time-series implications of our model with imperfect information with respect to the perfect-information counterfactual. The blue line shows the average reputation premium Υ (as a share of spreads) across paths in which at least one message $m = L$ was realized during 2007.Q1-2008.Q1 (our baseline sample period in the empirical analysis). The gray area shows a 90% confidence interval across *all* of the realized $\{m_t^i\}_{t=1}^T$ paths. The dashed black line shows the reputation premium under perfect information.

There are two main takeaways from the figure. First, the model with imperfect information delivers a more gradual increase in the reputation premium. Under perfect information, the reputation premium jumps upon the type switch (2007.Q1) and is significantly higher for the

Figure C.8. Model Simulations

(A) Path for Output (B) Reputation Premium (as % of spreads)



Notes: Panel (A) shows Argentina's log-linear cycle of GDP for the period 2006:Q1-2012:Q4. Panel (B) shows the reputation premium as a fraction of spreads. The solid blue line shows the case in which at least one $m = L$ was realized during 2007.Q1-2008.Q1. The gray area shows a 90% confidence interval across all $\{m_t^i\}_{t=1}^T$ paths. The black dashed line shows the perfect-information case.

first 5 quarters. Only after 6 quarters, there are $\{m_t^i\}_{t=1}^T$ paths within the 90% CI such that $\zeta \simeq 0$, and thus the reputation premium of our model almost coincides with the one under perfect information.[57] Second, although the reported CI bands exhibit some variation across the $\{m_t^i\}_{t=1}^T$ paths, the figure shows that across all of the paths in the 90% CI, Argentina's reputation premium can consistently account for more than a quarter of its spreads during the peak of the Great Financial Crisis.

C.6. *Solution Method*

We use a global solution method to solve the quantitative model described in Section 2. The state of the economy is (y, b, ζ). We discretize the output process y using Tauchen's method. We choose 15 gridpoints for y, 48 for b, and 15 for ζ. Gridpoints for the ζ grid are evenly spaced in the $[0, 1]$ range. We use 15 evenly spaced points for $b \in [0, 0.40]$ and 33 points for $b \in (0.4, 1.15]$. We use more points in the latter because the pricing kernel exhibits larger nonlinearities in that range. The steps of the algorithm are as follows:

[57] Small differences in the reputation premium still arise due to differences in the bond policies across the two counterfactuals.

(1) We start with a guess for the value functions $W_j(y, b, \zeta)$ for $j = \{C, S\}$. We also guess the lenders' conjecture $d^\star_j(y, b, \zeta)$ and $\tilde{\Pi}^\star_j(y, b, \tilde{\zeta})$, and the bond-pricing kernel $q(y, b', \zeta')$.

(2) At stage 1, if the government is not currently in default, the state of the economy is $(y, b, \tilde{\zeta})$, where $\tilde{\zeta} = \tilde{\zeta}(d = 0, \zeta, d^\star_C, d^\star_S)$—as shown in Equation (3). Taking $\tilde{\Pi}^\star_j(y, b, \tilde{\zeta})$ as given, for each message $m = \{L, NL\}$, we compute the lenders' posterior $\hat{\zeta}(m) = \hat{\zeta}(m, \tilde{\zeta}, \tilde{\Pi}^\star_C, \tilde{\Pi}^\star_S)$ based on Equation (4).

(3) Based on the guesses of step (1) and the updated posteriors of step (2), we can then solve for the optimal bond policy, $b^{\star\prime}(y, b, \tilde{\zeta})$—as described in Equation (A.4). To this end, we use a simple bisection algorithm (Brent's method) and we linearly interpolate the value functions and bond prices when evaluating off-grid points.

(4) Taking as given the solution for $b^\star(y, b, \tilde{\zeta})$, we solve for the S-type $\tilde{\pi}(y, b, \tilde{\zeta})$ policy, as described in Equation (A.6), where $V_j(y, b, \tilde{\zeta})$ is given by Equation (A.5). We use the same bisection algorithm of step (3). We then use $V_j(y, b, \tilde{\zeta})$ to compute $W^R_j(y, b, \zeta)$, following Equations (A.6) and (A.7).

(5) We compute the value function for the case in which the government defaults in the current period, $W^D_j\left(y, \tilde{\zeta}\right)$—as given by equation (A.2). We also compute the value function for the case in which the government is already in default, $\tilde{W}^D_j(y, \zeta)$—as shown in Equation (A.3).

(6) At stage 0, we solve for the government's optimal default choice—as shown in Equation (A.1). We then update our guess for $W_j(y, b, \zeta)$. Similarly to Chatterjee and Eyigungor (2012), we convexify the default decision in order to achieve convergence. In particular, we assume that in each period, the government's value function $W^D_j(\cdot)$ is subject to an i.i.d. shock $\epsilon_w \sim \mathcal{N}(1, \sigma_w)$ so that the government defaults if $W^R_j(\cdot) < W^D_j(\cdot) \times \epsilon_w$. We choose σ_w small enough ($\sigma_w = 0.0015$) so that the convexified solution does not significantly differ from the "true" solution of the model. Let $d_j(y, b, \zeta)$ denote the optimal default choice.

(7) Taking as given the conjectures $d^\star_j(y, b, \zeta)$ and $\tilde{\Pi}^\star_j(y, b, \tilde{\zeta})$, we update the bond-pricing kernel $q(y, b', \zeta')$ according to Equation (A.8).

(8) We update the guesses for the lenders' conjectures, $d^\star_j(y, b, \zeta)$ and $\tilde{\Pi}^\star_j(y, b, \tilde{\zeta})$, based on the updated solutions for $d_j(y, b, \zeta)$ and $\tilde{\pi}_j(y, b, \tilde{\zeta})$ (from steps 4 and 6).

(9) We iterate over the previous steps until convergence of the value function, conjectures, and bond-pricing kernel.