The Effect of Fiscal Constraint on Inflation: A Case Study From Argentina

Acknowledgements

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1 Introduction

Inflation has long interested economists as an object of study. From Cagan (1956), to the current discussions on the merits of targeting a 4% inflation rate, the field has concerned itself extensively with understanding its manifestations, its causes and its consequences. This paper contributes to the literature on the determinants of inflation by empirically estimating the effect of restricting financing options on the price level. I conduct the analysis in the setting of Argentina from 2012-2014 and find no significant results: I am unable to reject the null hypothesis of no effect of fiscal constraint on inflation.

Much theoretical work has linked fiscal behavior to inflation. Notably, Sargent and Wallace (1983) and Sargent (1992) describe how the inflation tax adjusts to fiscal conditions. In contrast, Sims (1994) and Woodford (1995) describe how fiscal policy adjusts to reconcile a predetermined price level. In *Recursive Macroeconomics*, Ljungqvist and Sargent summarize these and eight other monetary doctrines to show how fiscal and monetary policy interact with inflation; I extend their setup to the context of Argentina (Section 2.2) and assess its predictions (Sections 4 and 5). Easterly and Fischer (1990), and Bruno and Fischer (1990), amongst others, precede me in studying the relationship between fiscal constraint, seigniorage, and prices.

This literature has faced a key difficulty in the empirical estimation of this question: there is simultaneous causality because while fiscal constraint may lead to inflation, the reverse is also true. This thesis studies the first direction, which can be simply summarized as follows: the inability to issue debt can lead a government to rely on seigniorage to finance its deficit. In classical models—like those that exhibit 'money neutrality'—this increase of the money supply will result in a one-for-one increase in the price level, i.e. inflation. Indeed, through rational expectations, agents may foresee this action in advance and increase their inflationary expectations; this increases the price level today as producers raise the prices of their output in anticipation of future higher costs of inputs. Another channel operates through the exchange rate: fewer issues of external debt reduces the ratio of foreign to local currency, causing the currency to depreciate. The result is a higher domestic price level as the price of traded goods increases to match the depreciation; the price of non-traded goods may also increase if firms try to preserve the dollar value of their markups. Furthermore, there may be feedbacks between the two channels. For example, seigniorage may affect depreciation directly if the increase in the local money supply lowers the ratio of foreign to domestic currency and this leads to a run on the exchange rate.

However, the second direction hinders the empirical estimation of the first: inflation can also cause or exacerbate fiscal constraint. For example, in models with sticky wages, inflation reduces the popularity of a government as voters receive lower real wages. An unpopular government faces a higher cost of borrowing. Or, the government's fiscal position is weakened through the Olivera-Tanzi effect—tax obligations that are fixed in nominal terms are eroded with inflation when there is a long lag between collection and assessment (Olivera 1967; Tanzi 1977). It is true that in some cases inflation may improve the government's fiscal position. For example, the process of bracket creep—whereby inflation in wages places more taxpayers into higher tax brackets—may increase the real value of tax collection. Or, a government that is a net borrower may engineer bouts of inflation to erode the real value of nominal obligations such as long maturity government bonds (Calvo 1988). Either way, there is potentially a two-way street between fiscal constraint and inflation, and it is necessary to use some source of 'exogenous variation' to elucidate the effect of the former over the latter.

Argentina is a good testing ground for my hypothesis because it experienced a set of exogenous shocks to fiscal constraint and a high inflationary environment from 2012-2014.[1] Indeed, previous research by Hébert and Schreger (2016) demonstrates that the set of rulings on the case *NML Capital v. The Republic of Argentina* are reasonably exogenous shocks to Argentina's likelihood of default. They use the approach to estimate the cost of default using the impact on equity prices. In the context of their analysis they also produce an interesting finding: they estimate a positive and significant effect of Griesa's rulings on the price of the dollar. This leads me to think that a study of the broader hypothesis of the impact of these shocks on prices is worthwhile. Indeed, a favorable outcome for NML Capital implied a continuation of a conflict that had restricted Argentina's government from access to international capital markets since 2001 to fund its deficit. Hence, favorable/unfavorable rulings also impacted the likelihood of resolution of this conflict and the possibility of issuing debt abroad as a substitute for printing. If agents believe the government will be constrained in its financing options and consequently increase their inflationary expectations, the result should be higher inflation in the present.

To implement the empirical analysis, I make the key identifying assumption that

[1] Annual inflation during this period averaged 29.17% p.a. (Billion Prices Project).

the rulings affect prices only through their impact on fiscal constraint, i.e. there are no other channels through which these rulings impact inflation. I instrument for the likelihood of fiscal constraint through the spreads of Credit Default Swaps.

I implement two empirical strategies and find no significant effect of fiscal constraint on inflation. First, I conduct an instrumental variables analysis whereby I instrument for each ruling and conduct a standard two stage least squares regression. Second, I conduct a heteroscedasticity analysis in the style of Rigobon (2003) and Rigobon and Sack (2004). This method relies on the heteroscedasticity of the variance of the shock to the likelihood of fiscal constraint. I find no significant results in any iteration of these analyses.

Section 2 presents a discussion of the literature on inflationary costs, a theoretical analysis of the link between fiscal constraint and inflation, and historical evidence on the case *NML Capital v. The Republic of Argentina* as it relates to the question at hand. Section 3 presents the data and elaborates on the construction of the variables used in the empirical analysis. Section 4 implements the instrumental variables methodology and presents the results. Section 5 does the same for the heteroscedasticity based approach. Section 6 discusses the findings, the assumptions, and areas for improvement. Section 7 concludes.

2 Background

2.1 Literature

Generally, inflation is thought to be a pervasive, and in some cases costly, phenomenon and hence an important one to understand. It is so important that price stability is contemplated in most countries' central banks as a key component of their mandates to achieve stable growth (Fischer 1993). A target rate of 2% p.a. is generally considered the appropriate choice to accommodate growth while avoiding the costs associated with higher inflation rates. Furthermore, the process of disinflation is also well known to be costly to output (Fischer et al, 2002). Hence, most governments are wary of embarking on the inflation roller coaster.

The theoretical costs of anticipated and unanticipated inflation are summarized by Fischer and Modigliani (1978). For expected inflation, these are: shoeleather costs, menu costs, price dispersion and the inconvenience of a changing yardstick. Unexpected inflation has a more pernicious effect in that it informally induces redistribution of wealth between creditors and debtors through a process that is not seen as fair; the direction depends on whether inflation has been over/underestimated in the contracts between them.

However, in the literature it is hard to find a consensus on the nature and extent of these costs. Ball, Mankiw and Romer (1988) argue for an output-inflation tradeoff consistent with New Keynesian models of inflation. In line with this explanation, Ascari and Sbordone (2014) find that higher trend inflation destabilizes the economy and increases the cost of price dispersion. Further, Alvarez et al (2011) find that relative price dispersion indeed increases with high inflation in Argentina from 1989-1997. In contrast, Cavallo (2016) reports no change to the absolute size of price changes across different inflation regimes using scraped data.

Despite the lack of clear empirical consensus on the costs of inflation, the phenomenon is popularly considered problematic and symptomatic of a weak economy. Prominent thinkers and politicians have been known to comment on the harmful effects of inflation,[2] and surveys show it is at the forefront of the poor's minds (Easterly and Fischer 2001). Fischer et al (2002) find that since 1947, 20% of countries have experienced bouts of annual inflation exceeding 100%. Double digit inflation is also more common than many think. For example, the United States reported yearly inflation rates around 10% as recently as the 1970's. By private estimates, in the period 2012-2014, Argentina experienced yearly inflation rates of approximately 25% (Billion Prices Project). Hence, it is still important today to improve our understanding of the determinants of inflation. This work seeks to contribute to this endeavor.

2.2 A Model of Fiscal Weakness and Inflation

Inflation is defined as the change in the price level from one period to the next:

$$\pi_t = \frac{p_t}{p_{t-1}} \tag{1}$$

[2] "There is no subtler, no surer means of overturning the existing basis of society than to debauch the currency" John Maynard Keynes.
"[Inflation is] the cruelest tax" President Ronald Reagan.
"[Inflation is] public enemy number one" President Gerald Ford.
"[High inflation] undermines confidence and the ability of firms and households to make longer-term plans" Federal Reserve chairman Ben Bernanke.

where π_t denotes inflation and p_t is the price level in period t. Hence, to understand the determinants of inflation, it is necessary to understand what determines the price level in an economy. This section explores two models of price determination; the first is the Quantity Theory of Money and the second is an extension of Sargent and Ljungqvist's shopping time monetary economy model.

2.2.1 The Quantity Theory of Money

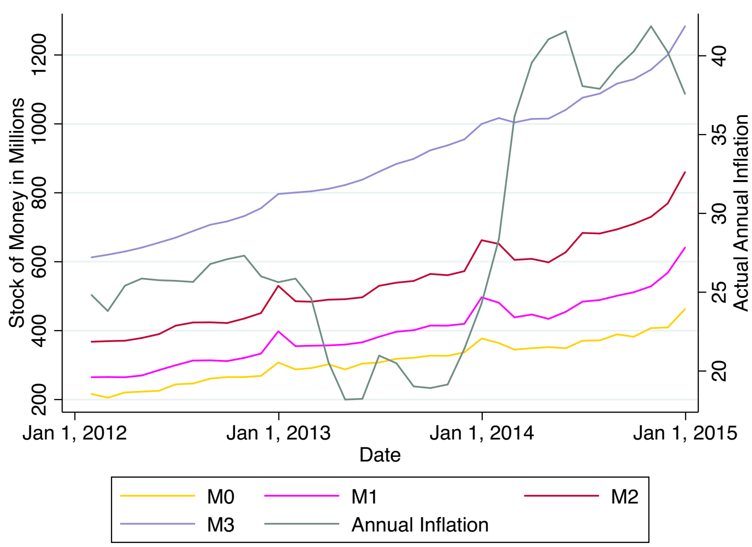
The canonical model to understand the determination of the price level is the quantity theory of money. In its simplest form, it is an identity:

$$MV = PT \tag{2}$$

where the left-hand side is made up of M, the stock of assets used for transactions, and V, the velocity at which it travels through an economy. The right-hand side can be estimated as the dollar value of output, or equivalently nominal GDP. In this case, the value of transactions, T, would be replaced by Y, the real value of output. Thus, if velocity V is assumed fixed and Y is determined by factors of production and the production technology, then the price level P is determined by M. The supply of money is determined by the central bank. This is typically thought of as a model of the 'long-run', a horizon over which monetary non-neutralities have washed out and GDP is determined by the supply-side of the economy. Indeed, over long time-frames, this theory seems consistent with the data: higher levels of M are associated with higher price levels.

Figure 1 shows the evolution of four measures of M (left axis) and annual inflation (right axis) for Argentina between 2012 and 2014. There are two moments when the printing of money accelerates, January of 2013 and January of 2014, which correspond to seasonal increases in the demand for money during end-of-year festivities, the payment of end-of-year bonuses,[3] and the summer vacation months. Annual inflation in 2012 is 25.64% p.a.. It drops to 24.3% p.a. in 2013 and rises dramatically to 37.58% p.a. in 2014. The reason for the drop in inflation in 2013 is a combination of the recession Argentina suffered that year and the implementation of price controls (Cavallo et al 2016). The subsequent increase in 2014 corresponds to the largest devaluation of the currency in 12 years, approximately 10%, during the final days of January.

Figure 1: Evolution of the Stock of Money and Inflation



However, this model is incomplete for the analysis at hand in two ways. First,

[3] In Argentina, the bonus paid at the end and the middle of the year is known as 'aguinaldo'. It is mandated by law.

it only contemplates how changes to the supply of money may lead to inflation. In contrast, this thesis focuses on how fiscal constraint causes inflation. The missing link, is between fiscal constraint and the supply of money. Secondly, this is only a model for the long run whereas this study is also concerned with the short run. The following extension of Sargent and Ljungqvist's model addresses these two concerns.

2.2.2 The Shopping Time Monetary Economy Model: An Extension

In this section I will extend Ljungqvist and Sargent's modified complete markets model with positive inconvertible currency value in order to work through the mechanics of how fiscal constraint may lead to inflation.[4] In this context, fiscal constraint is the inability to issue new debt (B in the model). This is representative of the problem Argentina faced in 2014 since the country was incurring the cost of repaying old issues of debt—it was servicing restructured bonds—but it was unable to issue new debt in international capital markets. The main reason it had been unable to do so since its original default in 2001 was the fact that it was still in conflict with the creditors who had not accepted the terms of the restructuring: of the $100 billion in debt that had been defaulted, $10 billion remained unresolved. At the heart of the case *NML Capital v. The Republic of Argentina* was the possibility of resolving the holdout problem and re-entering international capital markets.[5] Anecdotally, even when Argentina entered a selective default on July 30th, 2014, it still incurred the cost of servicing the restructured bonds since it forwarded payment of $539 million

[4]Please refer to chapter 26 of *Recursive Macroeconomics* for a full derivation of the model.

[5]It is true that Argentina was still able to issue debt domestically. The derivation in this section can easily be extended to account for this distinction by introducing a B* to represent foreign issues of debt. However, the mechanics work out the same.

to its trustee, the Bank of New York Mellon (BONY), which the latter retained until the full resolution of the conflict in 2016.

Basic Set Up

Ljungqvist and Sargent's model consists of an endowment economy with no uncertainty and a single good of constant amount $y > 0$ each period $t \geq 0$. There are two players: households and the government.

Households solve:

$$\sum \beta^t u(c_t, l_t) \tag{3}$$

Subject to,

$$s_t = H(c_t, \frac{m_{t+1}}{p_t}) \tag{4}$$

$$s_t + l_t = 1 \tag{5}$$

$$c_t + \frac{b_{t+1}}{R_t} + \frac{m_{t+1}}{p_t} = y - \tau_t + b_t + \frac{m_t}{p_t} \tag{6}$$

Where utility is increasing in c_t, consumption, and l_t, leisure. m_t are money balances such that $\frac{m_t}{p_t}$ are holdings of real money balances. s_t is shopping time, where $H_c > 0, H_{\frac{m}{p}} < 0$. τ are taxes and b is the real value of government bond holdings that mature at the beginning of t. R_t is the real interest rate.

The government wishes to finance the spending stream given by $\{g\}_{t=0}^{\infty}$, and is held to the budget constraint:

$$g_t = \tau_t + \frac{B_{t+1}}{R_t} - B_t + \frac{M_{t+1} - M_t}{p_t} \tag{7}$$

Where M_t is the stock of currency that the government has issued at the beginning of period t. B_t is government indebtedness to the private sector maturing at the beginning of period t.

Long Run

Equilibrium is defined by a price system $\{g_t, \tau_t\}_{t=0}^{\infty}$, a consumption sequence $\{c_t\}_{t=0}^{\infty}$, a debt sequence $\{B_t\}_{t=0}^{\infty}$, and a money supply sequence $\{M_t\}_{t=0}^{\infty}$, for which the following are true:

a) the household's optimum problem is solved with $b_t = B_t$ and $m_t = M_t$

b) the government's budget constraint is satisfied for all $t \geq 0$

c) $c_t + g_t = y$

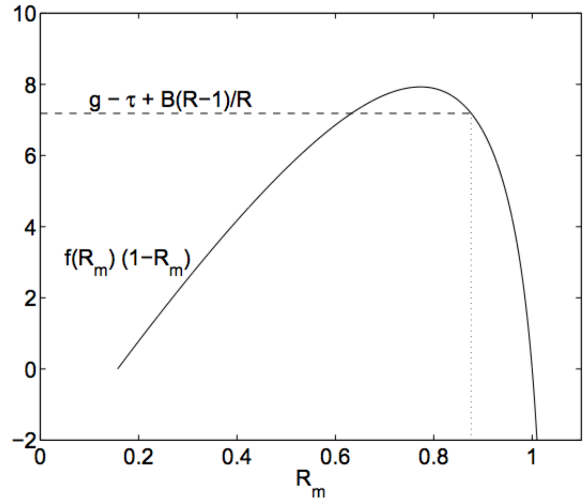
These conditions define the stationary rate of return on currency R_m [6] as follows:

$$g - \tau + \frac{B(R-1)}{R} = f(R_m)(1 - R_m),\ \forall t \geq 1 \tag{8}$$

Where the left-hand side of equation (8) represents the gross of interest government deficit and the right-hand side is the rate of seigniorage revenues from printing currency.

[6] $R_m = \frac{p_t}{p_{t+1}}$, and is the inverse of inflation $\pi_t = \frac{p_{t+1}}{p_t}$

Figure 2: Determination of the Stationary Rate of Return on Currency, R_m



The stationary rate of return on currency, R_m, is determined by the intersection between the stationary gross of interest deficit $g - \tau + B(1-R)/R$ and the stationary seigniorage $f(R_m)(1-R_m)$

Initial Date t=0

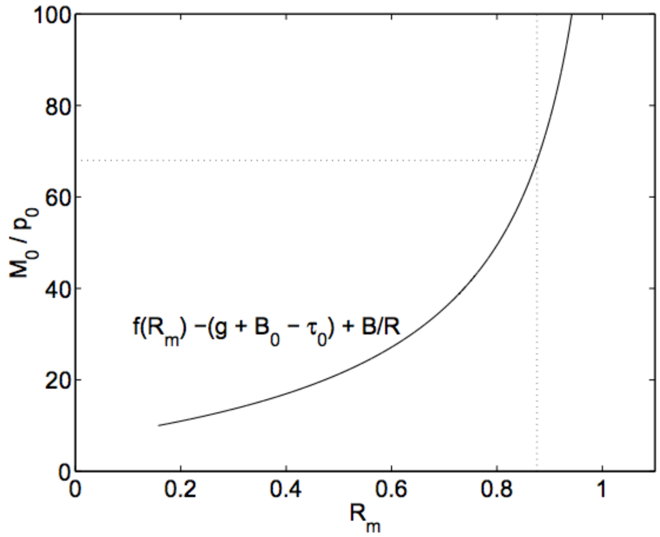
At $t = 0$, replace $\frac{M_1}{p_0} = f(R_m)$ to get:

$$\frac{M_0}{p_0} = f(R_m) - (g + B_0 - \tau_0) + \frac{B}{R} \tag{9}$$

Using equations (8) and (9), Ljungqvist and Sargent are able to study the effects of different monetary and fiscal doctrines on inflation. The dynamics of these equations can be seen visually in Figures 2 and 3.[7] The behavior of the economy is determined recursively; first, estimate what will happen in the long run (Figure 2). Then, with this value of R_m and using the fact that M_0 is given, determine the initial price level p_0.

[7] These are the same figures that can be found in *Recursive Macroeconomics*.

Figure 3: Determination of the Initial Price Level



Given R_m, the real value of initial money balances M_0/p_0 is determined by $f(R_m) - (g + B_0 - \tau_0) + B/R$. So, the initial price level p_0 is determined because M_0 is given

The effect of Fiscal Constraint in this model

Now, I extend this set up to understand the effects of fiscal constraint (FC) on inflation. Recall, fiscal constraint is interpreted as the inability to issue new debt B in the presence of a deficit. It is reasonable to assume that a government that receives bad news regarding debt will adjust its behavior e.g. lower spending, new ways to have people hold higher money balances or lie with statistics. However, a reasonable model where the optimal level of these variables is interior means that when the shock to debt hits, there will be more printing. Thus, for the sake of simplicity in this discussion, I take other fiscal variables such as taxes τ and spending g as fixed.

In the government's budget constraint, this looks like:

$$g_t = \tau_t - B_t + \frac{M_{t+1} - M_t}{p_t} \tag{10}$$

Notice the absence of the term $\frac{B_{t+1}}{R_t}$ in equation (10) as compared to equation (7) and recall that τ and g are pre-determined. Then, the long run equilibrium is given by:

$$g - \tau + B = f(R_m^{FC})(1 - R_m^{FC}),\ \forall t \geq 1 \tag{11}$$

Again, notice the absence of the term $-\frac{B}{R}$ in equation (11) as compared to equation (8). In a world of fiscal constraint, gross of interest government deficit is higher. This means that there is an upward shift in the gross of interest government deficit curve in Figure 2, which implies $R_m^{FC} < R_m$. A lower rate of return on currency implies a higher rate of inflation in the long run.

In the short run, the equation of the curve depicted in Figure 3 becomes:

$$\frac{M_0}{p_0^{FC}} = f(R_m^{FC}) - (g + B_0 - \tau_0) \tag{12}$$

where equation (12) differs from equation (9) by the absence of $\frac{B}{R}$. This translates to a downward shift of the curve as compared to the world of no fiscal constraint. Hence $\frac{M_0}{p_0^{FC}} < \frac{M_0}{p_0}$, which implies $p_0^{FC} > p_0$: there is also an increase in the short run price level when a government is fiscally constrained.

Thus, this model explains how fiscal constraint—the inability to issue debt—can lead to an increase in the price level both in the long and short runs. An important price not directly addressed in this model is that of the dollar. In developing countries, such as Argentina, the price of the dollar is also an important point of reference for the local price system[8]. A more complete model would do well to contemplate this factor.

2.3 Historical Context: ***NML Capital V. The Republic of Argentina***

In this section I review the history of Argentina's debt saga as it pertains to this analysis. This is useful in order to understand some of the assumptions made in the paper, such as the fact that it was unlikely Argentina would settle with the holdouts before January 2015, and to provide context for the interpretation of the empirical results.

On December 23rd, 2001, amidst a severe economic crisis that would result in an 11% drop in GDP, the Argentine government defaulted on $102.5 billion of external debt. This included loans from organizations such as the IMF, the World Bank and the Paris Club, as well as bonds held by private creditors. There were at least 152 different types of bonds denominated in 6 currencies and 8 jurisdictions (The Economist, 2005). After two debt restructurings, lump sum payments, and extensive litigation, the saga came to an end in 2016; in April, Argentina finally regained access to international capital markets.

In 2006 Argentina settled with the IMF in a total lump sum payment of $9.5 bil-

[8]One reason I do not focus on the price of the dollar in my analysis is that Hébert and Schreger (2016) already document an effect in this respect. My focus aims to extend this finding to the general price level.

lion. Despite the absence of any renegotiation, the deal was popular as it represented independence from the organization blamed for triggering the 2001 crisis. Thus, the Argentine government framed its fraught relationship with its creditors as an act of resistance; this fact is important to keep in mind when considering Argentina's behavior. Similarly, in May 2014, Argentina settled its debt with the Paris Club for $9.7 billion.

In contrast, private creditors were offered 35 cents on the dollar for their bonds. In 2005, after extensive and unfruitful negotiations, Argentina made its first such proposition to bondholders. The result: a 70% acceptance rate. The exchange offered three types of bonds issued across four currencies and four jurisdictions. $62.3 billion of the $81.8 billion owed in principal were exchanged for the new bond issues (Hornbeck 2013). The acceptance rate, which some have argued was high for such a drastic haircut, can be explained by two features of the new debt contracts. Firstly, the Argentine legislation passed a *'Lock Law'* which forbade the Argentine government from re-opening the exchange offer in the future and permanently suspended payments to bonds that were not exchanged. Secondly, the *'Rights Upon Future Offers (RUFO)'* clause, which expired in January of 2015, guaranteed to those who participated in the exchange that if the government voluntarily offered any creditor better terms, they would have a right to those terms as well. In 2010, Argentine legislation passed a temporary suspension of the *Lock Law* and the bond exchange was reopened under the same terms as in 2005. In total, 91.3% of bondholders accepted the exchange.

Argentina normalized payments to all restructured bonds and was declared out of default by S&P in 2005 as it upgraded the country's long term foreign currency credit rating to B- (MercoPress 2005). Subsequently the market began trading Credit Default Swaps on Argentina's restructured bonds. However, Argentina did not regain access to international capital markets until 2016. Instead, it benefited from high commodity prices, a rebounding economy and nontraditional self-financing methods. For example, in 2008, the government nationalized private pension funds for an estimated total of $25 billion (Partlow 2008).

The two restructuring deals left $11.2 billion in the hands of holdouts, of which $6.8 billion was in principal (Hornbeck 2013). Several of the holdouts held bonds that were under U.S. jurisdiction. Most importantly, this included NML Capital—a subsidiary of Elliot Management that specializes in distressed sovereign debt-who had purchased these bonds on the secondary market for 20 cents on the dollar in the early 2000's. In 2009, this group initiated litigation for their holdings of $1.3 billion. Their litigation was successful.

On December 7th, 2011, federal Judge Thomas Griesa, of the Second District Court, ruled that Argentina had violated the *Pari Passu* clause, which states that a debtor cannot give any creditor preferential treatment. Judge Griesa ordered that Argentina make ratable payments to the holdouts whenever it made a payment to the exchange bondholders. NML Capital argued that in paying the restructured bonds, Argentina was not respecting the equal treatment requirement. In contrast, Argentina argued that since the creditors who had participated in the restructuring held new issues of bonds, they should not be included in the *pari passu* consideration of the bonds held by the holdouts.

Even though the first ruling on this case occurred on December 7th, 2011, many rulings ensued as the case drew out for two years, traveled through three courts,[9] and gained international notoriety. Argentina disregarded the courts' 'ratable payments' order on the basis that its hands were tied by the RUFO clause until January 2015—until then it could not negotiate new terms without offering them to everyone else. In response, the holdouts filed motions to seize Argentine assets. The court ordered that the two parties negotiate and finally that no financial intermediaries could forward payments made to the restructured bondholders. Throughout the dispute, both sides presented various appeals. Judge Griesa's injunction was finally confirmed and put in place on June 16th, 2014 when the Supreme Court denied Argentina's second appeal.

On June 30th when the Argentine government transferred $539 million to the Bank of New York (BONY) to pay the exchanged bonds, Judge Griesa ordered that the funds be frozen as they violated his injunction. Thus, Argentina missed payments due on restructured bonds and following a month-long grace period, S&P declared Argentina had entered a 'selective default' on July 30th (Bloomberg 2014). The following day, the International Swaps and Derivatives Association decreed that credit default swaps on exchanged bonds would be settled.

Some features of this default are truly unique and merit mention, especially because they undermine the external validity of this work. Firstly, despite S&P's favorable rating as of 2005, in some sense the country had not resolved the 2001 default at the time of the court case. The trial itself was a result of the default and the

[9] The three courts were: United States District Court for the Southern District of New York, the United States Court of Appeals for the Second Circuit, and the United States Supreme Court.

country had been shut out of international capital markets since then. Hence, in 2014 Argentina defaulted while it was already bearing some of the costs from a previous default. Secondly, Argentina did have a willingness, and capacity, to pay its restructured bondholders. Indeed, bonds paid through national clearinghouses continued to receive payment; the price of these bonds increased following the default. Nevertheless, the events of 2014 are consistent with the interpretation that they are a negative shock to the Argentine government's fiscal position, and hence constitute an appropriate setting for this analysis.

Ultimately the situation remained in deadlock until the election of a new president in November of 2015. As part of his platform, Mauricio Macri promised to resolve the debt issue and return to international capital markets. He struck a deal with the holdouts in February of 2016 and the country was finally declared to be out of default in May. All defaulted debt was resolved by November of 2016 (MercoPress 2016). In April, Argentina 'triumphantly' returned to capital markets by issuing $16.5 billion in bonds (Hartley 2016).

3 Data

Primarily three data sets will be used in this analysis. The first is a proxy for the degree of fiscal constraint, the second is the set of events, and the third is a measure of inflation. The time frame for these data sets is from January 3rd, 2012, an arbitrary start date before the first event, to July 29th, 2014, the day before Argentina's 'selective default'.

3.1 The Events

Generally, I utilize Hébert and Schreger's (2016) set of 15 events. Each event is a ruling in the court-case *NML Capital v. The Republic of Argentina* and range from November 27th, 2012 to June 27th, 2014.[10] The authors discard rulings that coincide with public announcements by the government to avoid events that may violate their exclusion restriction. In their case, the concern is that the returns of stocks on these dates are impacted by something other than the probability of default. I respect this choice inasmuch as I also wish to identify the pure effect of fiscal constraint on inflation, i.e. I wish to avoid events where inflation may be responding to governmental announcements that in turn are endogenous to fundamental changes in the economy. In these cases, my coefficient would pick up an endogenous effect. Discarded rulings are classified as 'excluded' events and are removed from the sample set entirely.

The instances where I depart from Hébert and Schreger's set of 15 events depend on the type of analysis and the window of inflation being analyzed. For example, when I analyze daily inflation, I may consider all the events as established by Hébert and Schreger. However, when I analyze weekly inflation, events that are less than a week apart are problematic since the measure of inflation corresponding to each will overlap with the other. This means that the measure of inflation will be serially correlated. There is no standard satisfactory way of dealing with this problem. So, where necessary, I proceed by including only the latter of two (or more) unsuitably

[10]See appendix B for a detailed description of the events. See also Hébert and Schreger (2016) for a detailed discussion of each event and their rationale for inclusion/exclusion.

contiguous events. I gain economic precision in my estimates—both because I resolve the problem of serial correlation and because I take longer time windows which are more realistic in terms of allowing for inflation to respond to the shocks—but I lose statistical power by reducing the set of exogenous shocks. Any departure from the base set of events will be specified along with the results of the pertinent analysis.

I also construct a set of non-event days in order to provide a baseline over which to analyze the effect of the shocks. Non-event days are suitably far away from both included and excluded events and are non-overlapping. Naturally, as with the events, the choice of non-events varies based on the window of inflation under consideration. When analyzing daily inflation, I may for example consider both April 1st, 2014 and April 8th, 2014 as non-event days. However, when considering monthly inflation, I must choose between them to avoid serial correlation in the measure of monthly inflation. This need to drop observations explains why, even in the analysis of daily inflation, the number of observations is on the order of 200 instead of the order of 900 which one would expect over the span of three years.

3.2 The Likelihood of Fiscal Constraint

I proxy for the degree of fiscal constraint with the risk-neutral market-implied probability of default because there is a direct relationship between the two through the outcome of *NML Capital v. The Republic of Argentina.*

The risk-neutral market-implied probability of default is measured using Credit Default Swap (CDS) spreads that reference restructured bonds. Credit Default Swaps are financial instruments, defined in reference to a specific series of bonds, that allow investors to hedge against the risk of credit events of these bonds. What constitutes a credit event is determined by the International Swaps and Derivatives Association (ISDA) and can be any of the following: bankruptcy, obligation acceleration, obligation default, failure to pay, repudiation/moratorium and restructuring. The buyer of a Credit Default Swap must make periodic payments to the seller until a pre-agreed date. If, during this time a credit event—as defined by ISDA—occurs, the seller must pay the buyer the difference between the recovered value of the bond at auction, also known as the recovery rate, and the original value of the bond (Investopedia 2003).

In this case, since the CDS referenced the restructured bonds, the outcome of *NML Capital v. The Republic of Argentina* had a direct bearing on whether they would be settled or not. In particular, taking into account Argentina's aforementioned belligerent attitude, and the fact that the RUFO clause had not yet expired, if the outcome of the case was unfavorable for Argentina then it was unlikely the country would settle with the holdouts. If this occurred, Judge Griesa's injunction meant that payment to the restructured bonds would be interrupted—indeed this is what we know occurred. This in turn would trigger CDS settlement since interruption of payment constitutes a credit event, namely 'failure to pay'. On the contrary, a favorable outcome for Argentina meant a resolution of the holdout conflict without interruption of payment to the already restructured bonds and hence no CDS settlement. Thus, the likelihood of default as measured by these CDS spreads is directly impacted by the outcome of the case. This is also the reasoning made in Hébert and Schreger (2016).

The likelihood of fiscal constraint was also dependent on the outcome of *NML Capital v. The Republic of Argentina.* Similarly as above, if the outcome for Argentina was unfavorable, it was unlikely the country would settle with the holdouts. This meant a continuation of the holdout conflict, which in turn meant further exclusion from international credit markets i.e. further fiscal constraint. On the other hand, a favorable outcome for Argentina implied a resolution of the holdout conflict and the possibility of re-entry to international capital markets.

Hence, the common dependence on the outcome of *NML Capital v. The Republic of Argentina* makes the risk-neutral market-implied probability of default a good measures of the extent to which the government would be constrained in its financing options. White (2013) models the credit risk of bonds using the time varying hazard rate of default. From this model, it is possible to extract the so-called risk-neutral probability of default. This is the value I use as a measure of Argentine fiscal constraint in my analysis.

Thus, I calculate the risk-neutral probability of default using the credit triangle relationship as exposed by White (2013). I take Bloomberg quotes on the end-of-day spread of CDS until July 29th, 2014. Pricing before this date is questionable given the low transaction liquidity of these instruments, but it is the best high frequency measurement of the extent of Argentine fiscal constraint that is available to me. Pricing after this date is even more untrustworthy since the market was affected by the occurrence of a credit event and subsequent settlement of the instruments. In my empirical estimates I use the spread of the 5-year credit default swap, so the outcome should be interpreted as the risk-neutral probability that the Argentine government will default in the next 5 years.

Specifically, I implement a simplified version of the White (2013) model where I must assume that the premium leg is paid instantly and the hazard rate is equal to a constant λ.[11] In practice, these are strong assumptions. For example, professional financial forecasters such as Markit (2012) use time-varying hazard rates of default. Unfortunately, these calculations require data that was unavailable to me.

So, first I solve for the hazard rate, λ, using its relationship to the spread, S, as such:

$$S = (1 - R)\lambda \tag{13}$$

$$\lambda = \frac{S}{1 - R} \tag{14}$$

where R is the recovery rate. For the purposes of the current calculation I use a recovery rate of 39.5 that is the rate which was realized in the auction of CDS following the June 30th credit event.

Then, I approximate the risk neutral probability of default in the next five years by calculating:

$$Pr(D < 5Y) = 1 - \exp(-5\lambda) \tag{15}$$

Furthermore, the change in the probability of default is taken over a pre-determined window. Since I use it as a proxy for the change in the degree of fiscal constraint, I

[11]The 5-year hazard rate λ_t is the likelihood that Argentina will default 5 years from time t conditional on not having defaulted up until time t.

denominate it as ΔF_t and calculate it as:

$$\Delta F_t = Pr(D < 5Y)_t - Pr(D < 5Y)_{t-w} \tag{16}$$

where w represents the window size of the events. I calculate the change in the likelihood of fiscal constraint over 1-day event windows (w=1).[12] For example, the change in the likelihood of fiscal constraint associated with the ruling on Monday, June 16th 2014 should be the difference between the end-of-day default probability on June 16th and on June 15th. However, if an event falls on a Monday, I take the difference between the probability of default between Friday and Monday. Thus, in this particular case I take the difference between the likelihood of fiscal constraint on June 16th and subtract from it the likelihood of fiscal constraint on Friday, June 13th.

Figure 4 shows the evolution of the default probability between 2012 and 2014. Figure 5 shows the evolution of the change in the default probability between 2012 and 2014. Event dates are marked by red/blue lines representing unfavorable/favorable rulings for Argentina respectively. We would expect to see an increase in the default probability following unfavorable rulings and a decrease following favorable ones. On average the probability of default increases by 1.67 percentage points following the eleven unfavorable events and decreases by 4.36 percentage points on average following the four favorable events.[13]

[12]See Appendix C for results using 2-day event windows (w=2).

[13]See Appendix B for a detailed description of each event. Also see Hébert and Schreger (2016) for further information on rationale for inclusion/exclusion of all the rulings.

Figure 4: Evolution of the default probability

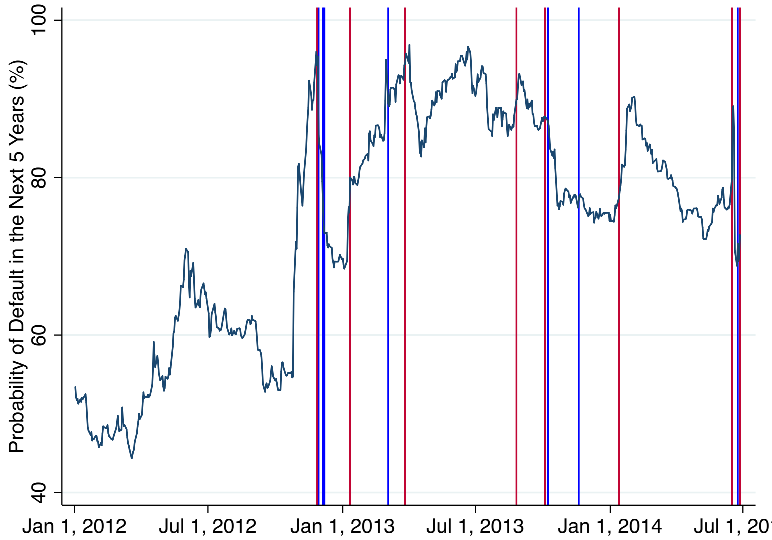


Figure 5: Evolution of the change in the default probability

3.3 Inflation

To estimate the occurrence of inflation I use data on both actual and expected prices. The data collection process and construction of indexes is described below.

3.3.1 Actual Inflation

This is the main measure of inflation used in the empirical analysis and is constructed using data from the Billion Prices Project (Cavallo and Rigobon 2016). The initiative was begun in 2007 by Alberto Cavallo as a response to the increase in untrustworthiness of official statistics and sought to provide an alternative to the official CPI published by the Argentine Government. The data set consists of scraped on-line prices from major retailers in Argentina. Since then they have expanded to other countries and smaller retailers in the region. The data has a daily frequency and is highly reliable since it is scraped directly from the retailer's website. There is a concern that retailers may report different values on-line than the one they offer off-line; however, studies have shown that this is not the case (Cavallo 2016).

In this paper, I use data from the supermarket COTO, one of the largest retailers in Argentina. They have over 120 branches, employ 19,000 people, and make $2.3 billion in annual revenue. Theirs is the most reliable and user-friendly data set as they have kept a consistent reporting method throughout this period. Although there are missing values when products are discontinued, the data set is still very large as it contains 38,335 unique products. Each data point consists of the row: id, product, date, url, price, control, sale as shown in Figure 6.

Figure 6: Example COTO Product Data Point

id	product	day	month	year	url	price	control	sale
69576	Fideos largos huevo DON VICENTE fettuccini paq 500 grm	16	2	2011	376	8.23	.	.

'id' is a unique numeric identification for each product type. 'product' is a qualitative description of the good including product name, brand and unit size. 'url' is an id for the webpage from which the product information was extracted. Usually, products of similar nature share the same url since COTO's on-line marketplace separates goods based on product category. For example, apples and bananas will share the same url since they both fall in the fruit category. The final two items are binary variables indicating whether the reported price is 'controlled' or on sale.

For the empirical analysis, I drop the ids that were controlled. This is an initiative by the Argentine government that seeks to make basic goods affordable by implementing price ceilings and therefore represents artificial price changes that could interfere with the analysis. So, it could happen 500gr of 'SanCor' butter is controlled while 500gr of 'La Serenísima' butter is not. Similarly, I drop the ids that had sales as they represent artificial deflation. The resulting data set has 21,983 unique id values.

Relative Prices

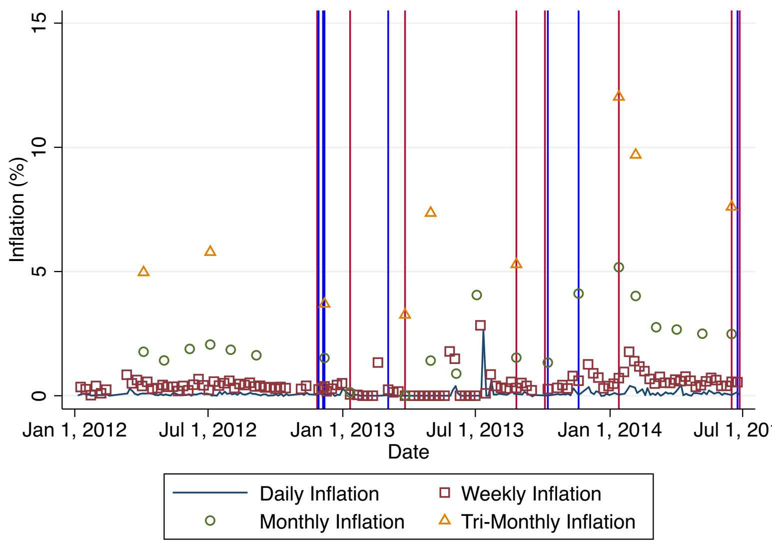
From this data set I calculate the inflation rate through the relative price change of each individual product as such:

$$\Pi_t = \exp[\frac{\sum_{i=1}^{n} \log(\pi_t^i)}{n}] \tag{17}$$

where Π_t is the aggregate rate of inflation at time t and π_t^i is the inflation rate of good i at time t for all $i \in (1, n)$. See appendix A for a full derivation of Equation 17. Note, to derive this price index I both assume Cobb-Douglas preferences in the underlying utility function, and I weight products equally in the basket. Since the focus of this work is general price movements, these are reasonable assumptions and methods.

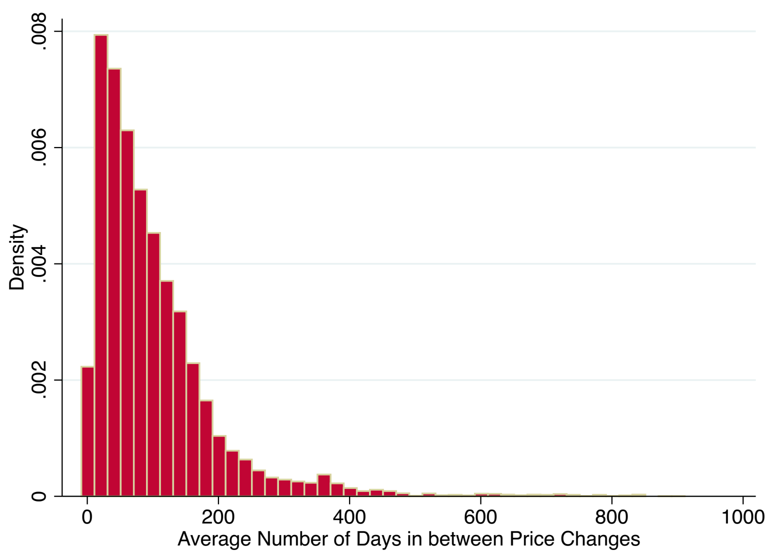
Furthermore, in the empirical analysis, I study the effect of the change in the likelihood of fiscal constraint on inflation over different time periods. This means that I estimate the effect of ΔF_t on daily, weekly, monthly and tri-monthly inflation (See Figure 7). As mentioned above, I make sure to take non-overlapping periods to avoid serial correlation. Thus, as I take larger windows, my sample size of observations gets smaller.

Figure 7: Evolution of Inflation



Inflation can occur through two channels, the frequency of price change and the absolute value of a price change. For example, take an Apple and a Banana that are both worth 10 pesos at $t = 1$. If the Apple's price increases by 1 peso on each of the following 10 days and the Banana's price increases by 10 pesos on the 10th day, then both experienced 100% inflation by $t = 10$. However, the Apple experienced a higher frequency of price change while the Banana experienced a higher absolute value of price change. In the sample, on average, prices changed every 100 days.[14] Figure 8 is a Histogram of the observed average price change. As can be observed, the distribution of price changes has a long right tail. This means there are a few products that change prices very infrequently. I do not exploit changes in the frequency of price changes as a measure of inflation in my analysis, although this would be an interesting strategy for further work. Instead, I use this observed frequency as a rule of thumb for when I should expect to see the effect of ΔF_t show up in measures of inflation. A frequency of 100 days in between price changes suggests that tri-monthly inflation would be the most likely measure to pick up the effect of changes to the likelihood of fiscal constraint for the majority of prices. I report the effect on several windows of inflation for the sake of completeness.

Figure 8: Frequency of Price Changes



[14]This value is calculated by taking ids whose price changed more than once in the sample. Then, for each of these id's I count the number of days between the first recorded price change in the sample and the last recorded price change in the sample. Then, I divided this window by the number of price changes minus one. Thus, the outcome should be interpreted as the average number of days between price changes. In this case, prices on average changed every 100 days.

Price Dispersion

Next, I analyze the effect of a change in the likelihood of fiscal constraint on different measures of price dispersion. The idea that prices will disperse when there is inflation is closely related to the menu cost model of inflation as it predicts that firms with low menu costs will rapidly update their posted prices to follow inflation while those with high menu costs may lag behind. The objective of this strategy is to take advantage of the disaggregated data provided by the Billion Prices Project and to provide a more complete analysis of price behavior in Argentina during this period.

The first measure of price dispersion builds on Alvarez et al (2015) who find that the dispersion of relative prices is an increasing function of inflation for high levels of inflation. Accordingly, to estimate whether fiscal constraint causes inflation, I estimate whether the dispersion of relative prices increases when the likelihood of fiscal constraint increases. This measure assumes that the inflationary shock hits the system in equilibrium where every firm is at its optimal price. Then, inflation generates dispersion as some firms can update their prices and some firms cannot.

Thus, I first calculate the cross-sectional variance of product specific inflation over different time windows as such (example shows daily inflation):

$$\text{Var}(\text{Daily } \pi_t) = \text{Var}\left(\frac{p_t}{p_{t-1}}\right) = \frac{\sum_{i=1}^{N}(\pi_{it} - \bar{\pi}_t)^2}{N} \tag{18}$$

Where $\pi_{it} = \frac{p_{it}}{p_{i(t-1)}}$ is the daily inflation of product i at time t. There are N products.

For this measure to be valid, I must also test whether the system was 'at rest'.

To do so I conduct an augmented Dickey-Fuller test for stationarity and find that for daily, weekly and monthly measures of relative price dispersion I can reject the null hypothesis that the system was not at rest at the 95% confidence level. Unfortunately, I cannot reject non-stationarity of variance in tri-monthly inflation. I report all the results in Section 4.2.1 for the sake of completeness.[15]

The second measure of price dispersion looks at how the change in the likelihood of fiscal constraint impacts the change in the cross-sectional variance of prices. I use a subset of the data that includes only the products that are present for the entire estimation period. The reason for this is that I am concerned with artificially skewing the results if some products enter and leave the sample since I compare the variance of price levels. This sample still includes 5,748 unique ids, so arguably I am not biasing the sample in other ways such as small sample size or an unrepresentative sample, such as only shampoo products. I construct the left hand side variable as such (example shows 1-day window):

$$\Delta_1 \text{Var}(p_t) = \frac{\text{Var}(p_t)}{\text{Var}(p_{t-1})} = \frac{\sum_{i=1}^{N}(p_{it} - \bar{p}_t)^2}{N} / \frac{\sum_{i=1}^{N}(p_{i(t-1)} - \bar{p}_{t-1})^2}{N} \tag{19}$$

Where p_{it} is the price level of product i at time t. There are N products.

3.3.2 Expected Inflation

A second measure of inflation that I use in my analysis is expected inflation. I recognize that inflation is a slow moving variable (in the previous section I found

[15] Note that to conduct the unit-root test I must address the missing values in the series generated by the fact that I cannot have overlapping data points. In effect, I 'close up the gaps' as per Ryan and Giles (1998).

that on average prices changed only every 100 days). Hence, it may take time for inflation to incorporate the information from a shock to the likelihood of fiscal constraint which means this data set is likely noisy. Variables that might incorporate this information faster are measures of expected inflation. I use two measures; the first are inflation-indexed bonds and the second is a professional forecast. Their strengths and weaknesses are discussed below.

Inflation-linked Bonds

In the time period 2012-2014 there were three series of domestically issued bonds whose coupons adjusted by the official report of inflation. The concern has already been raised that official inflation did not reflect actual inflation. The discrepancy between the two is reflected in Figures 9 and 10. However, the correlation between the change in the official inflation and the change in the actual inflation is 0.69 for annual inflation rates and 0.49 for monthly rates of inflation. These values may be low, but they are positive. Therefore, we may cautiously argue for the use of official inflation as a proxy for actual inflation rates. Hence, shocks to actual inflation should be reflected in official rates as well and will therefore be reflected in the pricing of inflation-indexed bonds.

Figure 9: Official vs. Actual annual inflation

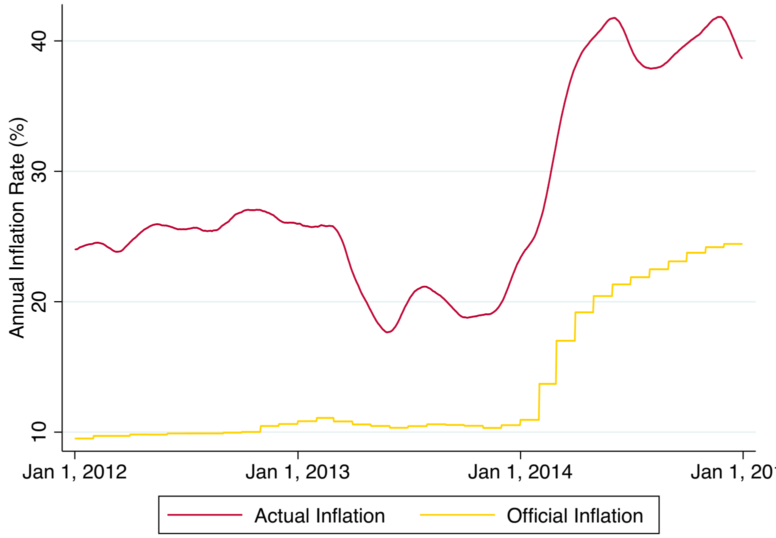
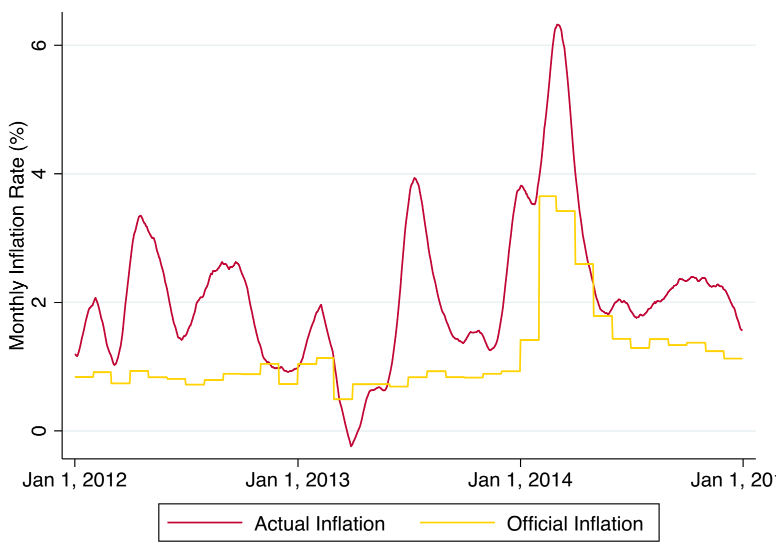


Figure 10: Official vs. Actual monthly inflation



If official rates of inflation react to shocks similarly to actual rates of inflation, then it is plausible to exploit the yields of inflation-indexed bonds to anticipate occurrences of inflation. The strength of this approach is that I will have a high frequency of observations, whereas the weakness is the noise created by the discrepancy between official and actual inflation.

Another drawback is the fact that the yields of these bonds were affected by the rulings through a channel other than their effect on inflation. Namely, if Argentina's fiscal position was strengthened/weakened by these rulings, then the countries ability to maintain payments would increase/decrease. Hence, there are two opposing effects to a negative fiscal shock: yields should decrease because of higher expected (official) inflation,[16] but they should increase because of the higher risk. This fact implies that any effect I find is an underestimate of the expected inflationary response to fiscal shocks.

The bonds that were current in this period, that were not part of the 2005/2010 restructuring, and that adjusted by official inflation were: Bocon 2016 (pr12), Bocon 2014 (pre9) and Bogar 2018 (nf18). For the first two bonds I use the last known yield downloaded from Bloomberg, while for Bogar 2018 I utilize last known price downloaded from Ravaonline, since this data set was more complete. Unfortunately, some of these bonds expired during the sample period. For each bond, I only consider events that occurred during the period for which they were current.

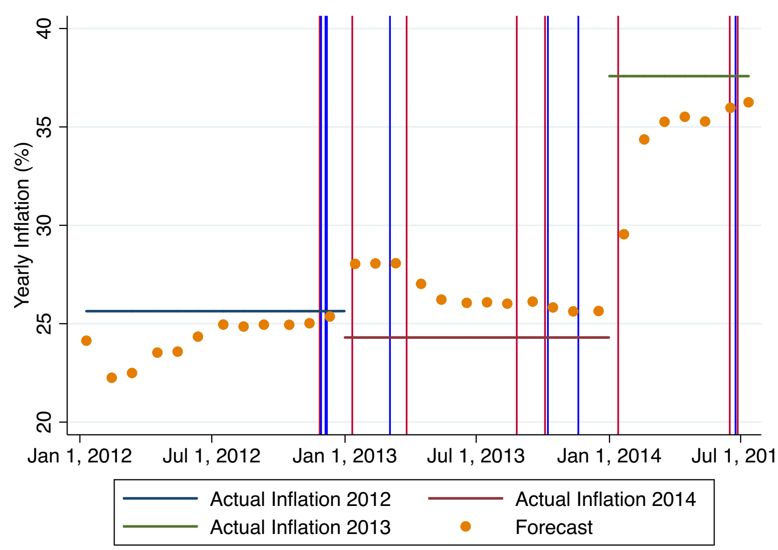
Since the bonds are financial instruments whose prices adjust immediately to contemplate all available information, I construct the dependent variable in these estimations as the change in either the yield or the price over the same 1-day window as the change in the likelihood of fiscal constraint.

[16]When a bond's coupon adjusts with inflation, the value of the coupon increases proportionally to inflation. Hence, with higher inflation the coupon's payments will be higher. A bond whose coupons are expected to increase will trade at a higher price implying a drop in yields.

Professional Forecast

I also estimate the effect of fiscal constraint on expected inflation using a professional forecast. The data set used here is published in the LatinFocus Consensus Forecast by FocusEconomics. It consists of monthly forecasts of yearly inflation for the current year collected on the second or third Tuesday of every month. So, for example, in January 2012 FocusEconomics collects and publishes predictions for the actual inflation that will occur in the year 2012. In February 2012 it publishes updated forecasts of this measure. Refer to Figure 11 for a graphical representation of the data. The fact that the data is collected on the second or third Tuesday of every month is important because of the timing of the rulings. Rulings that occur before said Tuesday will be contemplated in that month's report. However, rulings that occur in the same month but after this date will impact the forecasts reported in the following month.

Figure 11: Professionally Forecast Inflation



Thus, in order to conduct the analysis with this data set I match the timing of the rulings with the corresponding forecast they impact. Then, forecasts for which there is a corresponding ruling become 'event' months while the rest are 'nonevent' months. For each 'event' month I take the sum of the changes to fiscal constraint corresponding to each of the rulings in that month. So, for example, the October forecast of 2013 was impacted by the rulings on the 4th and the 8th of October, 2013. Thus, I take the change in the fiscal constraint probability corresponding to these two dates—0.00467 and -0.00589—and sum them to get the cumulative exogenous change to the probability of default over this period. For 'nonevent' months, I take the change in the probability of fiscal constraint on a random day within that reporting period. Finally, the dependent variable in this case is the change in the reported expected yearly inflation from one period to the next. Per the hypothesis, I expect to see an increase in inflationary expectations when the likelihood of fiscal constraint increases.

One of the drawbacks of this dataset is that forecasts made in the final months of the year become more precise simply because there is certainty over the inflation that has occurred thus far. Thus, the magnitude of the difference from one period to the next is inherently smaller. Inasmuch as forecasters become precise at the same rate each year, I can include time dummies in my estimations for each month to correct for this bias. A related drawback is the fact that these forecasts are not made on a rolling basis (i.e. the inflation that will occur in the following 12 months), but rather static to the year in which they are made. Finally, the frequency of this data set is low—expectations are only reported on a monthly basis. The benefits of using this dataset are that the forecasts refer to actual inflation (not official) and that they are done by professionals. Regarding the first, it is important to note that between 2007-2016 the national institute for statistics (INDEC) was intervened by the government and published unreliable statistics. Regarding the second point, studies have shown that households' perception of inflation is biased as compared to professionals' (Cavallo et al 2014).

4 Identification through Instrumental Variables

Instrumental Variables is a common strategy to resolve the issue of simultaneous causality bias. Originally this method was developed to estimate models of supply and demand. A canonical example is P.G. Wright (1928), who implemented instrumental variables to estimate the supply and demand schedules for flaxseed. The key to this methodology is to identify an 'instrument': a shock that is exogenous to the dependent variable but relevant to the independent variable. Angrist and Krueger (2001) review the evolution of the instrumental variables approach and highlight its effectiveness when properly implemented. Romer and Romer (1989) describe the process of identifying these instrumental variables as a "narrative approach" to econometrics since a certain amount of discretion on behalf of the researcher is necessary to establish the exogeneity of the instrument. However, if these assumptions are sound, then the analysis is meaningful. In the following section I will implement the instrumental variables approach to empirically answer the question of whether fiscal constraint leads to inflation.

4.1 Empirical Framework

I construct an instrumental variable, Z, using subsequent rulings on the case *NML Capital v. The Republic of Argentina*. I argue that Z_t is uncorrelated with Π_t since these rulings are based on legal interpretation of sovereign bond contracts. This is the first part of the identifying assumption: the legal rulings are uncorrelated with inflation. The impact of Z_t on Π_t thus occurs only through the effect of Z on the likelihood of fiscal constraint.

Furthermore, the rulings are relevant to the likelihood of fiscal constraint. This is the second part of the identifying assumption: the variable Z is correlated with the likelihood of default. If Argentina did well in the court case, it would be likely they would resolve the problem with the holdouts and re-enter international capital markets. This means they could issue new external debt and loosen their fiscal constraint. A more comfortable fiscal position means less inflation. On the contrary, an adverse outcome meant Griesa's injunction would be ratified and it was likely Argentina would miss its payments to the restructured bonds. The holdout conflict would continue and Argentina would remain fiscally constrained.

The instrumental variables approach proceeds in two stages. First, estimate ΔF_t using Z to isolate exogenous changes to ΔF_t. Then, use this estimated value of ΔF_t to estimate Π_t.

1st stage:

$$\Delta F_t = \alpha_0 + \alpha_1 Z_t + v_i \tag{20}$$

2nd stage:

$$\Pi_t = \beta_0 + \beta_1 \Delta \hat{F}_t + u_i \tag{21}$$

where α and β are the coefficients of the estimation and u and v are the error terms of the regression.

The Instrument

The instrument, Z, consists of individual dummy variables for each event date. For example, when estimating daily inflation I will have 15 dummy variables. However, when I estimate monthly inflation, I only have 8 dummy variables since I am forced to drop overlapping events in order to avoid serial correlation. The concern is that although the windows become more realistic in terms of the time it would take for inflation to react to a shock, the estimated coefficients may become biased due to the smaller sample size.

It is important to note that while the instruments arguably fulfill the two conditions for a valid instrument,[17] in the empirical analysis I find that some of the instruments are weak. This occurs when the F-test is less than 10. The most likely explanation is that the instruments are still exogenous,[18] but are not relevant. The consequence of having weak instruments is that the process of 2 stage least squares (2SLS) described will be biased towards OLS, and the usual standard errors and confidence intervals are not valid. Andrews and Stock (2005) evaluate existing methods for inference with weak instruments and suggest the use of Anderson and Rubin (1949) confidence intervals in these cases. So, wherever there are weak instruments, I also report these A-R confidence intervals.

[17] As laid out above, a valid instrument must be exogenous and relevant.

[18] Previous work, namely Hébert and Schreger (2016), has successfully exploited this fact.

4.2 Results

The following section presents the results for the instrumental variables analysis.

4.2.1 Actual Inflation

Relative Prices

Table 1 reports the effect of an increase in the likelihood of fiscal constraint on different measures of actual inflation as measured by relative prices.

Table 1: Results for Instrumental Variables Analysis Actual Inflation I

VARIABLES	(1) Daily Inflation	(2) Weekly Inflation	(3) Monthly Inflation	(4) Tri-Monthly Inflation
ΔF_t	-0.00184	0.0118	-0.106	0.618
	(0.00163)	(0.00859)	(0.209)	(0.477)
Constant	1.001***	1.005***	1.022***	1.063***
	(0.000104)	(0.000609)	(0.00267)	(0.0102)
Observations	644	115	21	9
IV F-Test	28122	7248.77	0.29	11.29

The change in the likelihood of fiscal constraint is calculated over a 1-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address the serial correlation in the data (Newey West, 1987). For weak instruments, i.e. (3) Monthly Inflation, Anderson Rubin confidence intervals are more appropriate (Anderson Rubin, 1949). In this case the A-R confidence interval for Monthly Inflation is the entire sample space.

All results are insignificant. However, it is interesting to note that while daily, weekly and monthly inflation decrease when the likelihood of fiscal constraint increases, tri-monthly inflation increases. This is consistent with the finding that prices change on average every 100 days. Hence, the effect of an increase in the likelihood of fiscal constraint should not be fully observed until approximately three months posterior to the date of the shock.

More specifically, a 10 percentage point increase in the likelihood of fiscal constraint causes a 6 percentage point increase in tri-monthly inflation. However, the smallest effect that can be ruled out with a 95% confidence interval is a 3 percentage point decrease and the largest is a 15 percentage point increase.

Unfortunately, the tri-monthly inflation rate is a very noisy measure due both to the large window of time it covers and the small sample size. Ultimately it is not possible to reject the null hypothesis of no effect of fiscal constraint on inflation.

Price Dispersion

Table 2 presents the results of the effect of a change in the likelihood of fiscal constraint on the cross-sectional variance of different measures of inflation.

Table 2: Results for Instrumental Variables Analysis of the Variance of Actual Inflation I

VARIABLES	(1) Var(Daily π)	(2) Var(Weekly π)	(3) Var(Monthly π)	(4) Var(Tri-Monthly π)
ΔF_t	0.00105	0.523	-0.0464	1.255***
	(0.000768)	(5.167)	(0.200)	(0.299)
Constant	0.000592***	0.764	0.0171***	0.0436***
	(8.80e-05)	(0.714)	(0.00231)	(0.00832)
Observations	214	115	21	9
IV F-Test	5267.47	7248.77	0.29	11.29

The change in the likelihood of fiscal constraint is calculated over a 1-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address serial correlation in the data (Newey West, 1987). However, for weak instruments, i.e. (3) Var(Monthly π), Anderson Rubin confidence intervals are more appropriate (Anderson Rubin, 1949). In this case the A-R confidence interval is the entire sample space.

Generally—the exception is monthly inflation—the results suggest that the cross-sectional variance of relative prices increases as a result of an increase in the likelihood of fiscal constraint. Most interestingly, column (4) states at the 99% confidence level that a 10 percentage point increase in the likelihood of fiscal constraint results in approximately a 12 percentage point increase in the relative price dispersion over a period of three months. Using price dispersion as a proxy for inflation, this result is consistent with the hypothesis that higher fiscal constraint leads to higher inflation. However, as mentioned in the data section, it is not possible to reject the null of no stationarity for this variable. Hence, this coefficient is meaningless.

Table 3 presents the results of the effect of a change in the likelihood of fiscal constraint on the change in the cross-sectional variance of the price level over different windows.

Table 3: Results for Instrumental Variables Analysis of the Change in the Cross-sectional Variance of Prices I

VARIABLES	(1) Δ_1Var(P_t)	(2) Δ_7Var(P_t)	(3) Δ_{30}Var(P_t)	(4) Δ_{90}Var(P_t)
ΔF_t	0.00876**	0.113	0.109	0.393
	(0.00438)	(0.0703)	(0.520)	(1.428)
Constant	1.001***	1.007***	1.034***	1.099***
	(0.000267)	(0.00159)	(0.00811)	(0.0205)
Observations	214	115	21	9
IV F-Test	5267.47	7248.77	0.29	11.29

The change in the likelihood of fiscal constraint is calculated over a 1-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address serial correlation in the data (Newey West, 1987). However, for weak instruments, i.e. (3) $\Delta_{30}Var(P_t)$, Anderson Rubin confidence intervals are more appropriate (Anderson Rubin, 1949). In this case the A-R confidence interval for both is the entire sample space.

The results suggest that indeed an increase in the likelihood of fiscal constraint causes the dispersion of the price level to increase over all windows. Although, surprisingly, the one day window is the only significant result, the 90-day window reports the largest effect. The result in column (1) is surprising based on my analysis of the frequency of price changes. Although the result is insignificant at the 95% confidence level, Column (4) suggests that a 10 percentage point increase in the likelihood of fiscal constraint causes a 3 percentage point increase in the dispersion of the price level.

4.2.2 Expected Inflation

Table 4 reports the effect of a change in the likelihood of fiscal constraint on expected inflation using the instrumental variables methodology for selected inflation-indexed bonds and a professional monthly forecast. Recall that due to data constraints, Bocon 2016 and Bocon 2014 reflect changes in yields, which implies that I expect a negative coefficient, while Bogar 2018 reflects changes in prices, which implies that I expect a positive coefficient under my hypothesis. Furthermore, the dependent variable in the fourth column is the month-to-month difference in inflation forecasts by professionals. Here, I expect a positive coefficient.

Table 4: Results for Instrumental Variables Analysis of Expected Inflation I

VARIABLES	(1) Bocon 2016	(2) Bocon 2014	(3) Bogar 2018	(4) Professional Forecast
ΔF	4.890**	2.205	-15.75***	-0.996
	(2.298)	(1.977)	(4.475)	(2.022)
Month Dummy				-0.0572
				(0.0676)
Constant	-0.0432	-0.0164	0.117***	0.559
	(0.0300)	(0.0426)	(0.0185)	(0.589)
Observations	399	307	567	30
IV F-Test	6002.89	4788.10	49793.52	81.98

The changes in the likelihood of fiscal constraint, the changes in the yield of Bocon 2016/Bocon 2014 and the changes in the price of Bogar 2018 are calculated over a 1-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address serial correlation in the data (Newey West, 1987).

Contrary to the hypothesis, and significant to the 95% and 99% confidence level respectively, the yield of Bocon 2016 bonds increased by 48 percentage points and the price of Bogar 2018 bonds fell by $1.5 Argentine pesos when there was a 10 percentage point increase in the likelihood of fiscal constraint. Also contrary to the hypothesis, although not significantly so, the yield of Bocon 2014 bonds increased and the levels of professionally forecast inflation fell in response to an increase in the likelihood of fiscal constraint. The behavior of the bonds suggests that the market is expecting (official) inflation to fall. It is possible that the higher risk of non-payment due to fiscal weakness is offsetting the effect of higher expected inflation as explained in the Data section.

5 Identification through Heteroscedasticity

An alternative method of addressing simultaneous causality bias was developed by Rigobon (2003) and Rigobon and Sack (2004). They exploit the heteroscedasticity of the shocks to their independent variable of interest to estimate its effect on the dependent variable. In order to apply this methodology, I make the assumption that the variance of fiscal shocks is higher on event days than on non-event days. The following section develops the empirical framework and presents the results.

5.1 Empirical Framework

The key reason for preferring this method is that it requires a weaker set of assumptions than the traditional OLS estimate to be unbiased. Namely, this methodology does not require the absence of all other common shocks on 'event days'. This may be a safe assumption if the event windows are very small. However, in this context, the necessity for larger windows suggests the heteroscedastic approach is more appropriate. The latter only requires the variance of the shock to ΔF_t be higher on event days than non-event days. Then, the distribution of observations would approach the inflation rate reaction schedule and identification is possible. Furthermore, the concern with the previous methodology regarding small sample size is mitigated here since previous papers have implemented this same methodology with 15 events (Hébert and Schreger 2016).

Recall, I wish to estimate α in the following set of simultaneous equations:

$$\Delta F_t = \beta \Pi_t + \gamma z_t + \varepsilon_t \tag{22}$$

$$\Pi_t = \alpha \Delta F_t + z_t + \eta_t \tag{23}$$

Where ΔF_t is the change in the likelihood that the government will be fiscally constrained, proxied by the change in the likelihood of default, and Π_t is a measure of inflation. z_t are common shocks to the likelihood of fiscal constraint and inflation, η_t are idiosyncratic shocks to inflation and ε_t are idiosyncratic shocks to the likelihood the government will be fiscally constrained. The idiosyncratic shocks are assumed to be serially uncorrelated, and uncorrelated with the common shock.

As stated before, the problem with estimating α directly is the simultaneity causality bias between ΔF_t and Π_t. So, following Rigobon and Sack (2004) I look at co-movements of the likelihood of fiscal constraint and inflation when the variance of the ε_t shock shifts. Specifically, I assume that the variance of the shock to fiscal constraint is higher on event days than on non-event days.

To estimate α, first solve for the reduced form of the system:

$$\Delta F_t = \frac{1}{1 - \beta\alpha}[(\beta + \gamma) z_t + \beta \eta_t + \varepsilon_t] \tag{24}$$

$$\Pi_t = \frac{1}{1 - \beta\alpha}[(1 + \alpha\gamma) z_t + \eta_t + \alpha \varepsilon_t] \tag{25}$$

The Variance-Covariance matrix of ΔF_t and Π_t is given by:

$$\Omega^i = \begin{bmatrix} Var(\Delta F_{it}) & Cov(\Delta F_{it}, \Pi_{it}) \\ . & Var(\Pi_{it}) \end{bmatrix} \tag{26}$$

where $i \in \{1, 2\}$ indicates an event or a non-event day respectively.

Thus,

$$\Omega^i = \frac{1}{(1-\beta\alpha)^2} \begin{bmatrix} (\beta+\gamma)^2\sigma^2_{z_i} + \beta^2\sigma^2_{\eta_i} + \sigma^2_{\varepsilon_i} & (\beta+\gamma)(1+\alpha\gamma)\sigma^2_{z_i} + \beta\sigma^2_{\eta_i} + \alpha\sigma^2_{\varepsilon_i} \\ . & (1+\alpha\gamma)^2\sigma^2_{z_i} + \sigma^2_{\eta_i} + \alpha^2\sigma^2_{\varepsilon_i} \end{bmatrix} \tag{27}$$

Hence,

$$\Delta\Omega = \Omega^1 - \Omega^2 = \frac{(\sigma^2_{\varepsilon_1} - \sigma^2_{\varepsilon_2})}{(1-\beta\alpha)^2} \begin{bmatrix} 1 & \alpha \\ \alpha & \alpha^2 \end{bmatrix} \tag{28}$$

Luckily, we can estimate $\Delta\hat{\Omega}$ from our data with two sets of dummy variables, one for event days and one for non-event days. This method of estimating α uses instrumental variable regression.

Thus:

$$\hat{\alpha} = \frac{\Delta\hat{\Omega}_{12}}{\Delta\hat{\Omega}_{11}} \tag{29}$$

Estimating directly from the sample, we obtain:

$$\hat{\alpha} = \frac{Cov(\Delta F_{1t}, \Pi_{1t}) - Cov(\Delta F_{2t}, \Pi_{2t})}{Var(\Delta F_{1t}) - Var(\Delta F_{2t})} \tag{30}$$

To calculate the 95% confidence interval of $\hat{\alpha}$, I implement the standard bootstrapping method with 1000 iterations. Results of this analysis are presented below.

5.2 Results

5.2.1 Actual Inflation

Relative Prices

Table 5 reports results using a window of 1-day for the change in the probability of fiscal constraint.

Table 5: Results for Heteroscedasticity Analysis of Actual Inflation I

	Daily Inflation	**Weekly Inflation**	**Monthly Inflation**	**Tri-Monthly Inflation**
ΔF_t	-.0008935	.0079702	-.1430634	6.171184
Bootstrap Standard Error	.0049167	.1473738	25.74613	102.6072
95% Confidence Interval	(-.011, .009)	(-.281, .297)	(-50.6, 50.3)	(-194.9, 207.3)
Events	15	11	8	5
Observations	214	117	21	9
Bootstrap Replications	1000	1000	999	967

The change in the likelihood of fiscal constraint is calculated over a 1-day window. All estimations are insignificant at the 95% confidence level.

As can be seen by the table, all results are insignificant given a 95% confidence interval. However, it is interesting to note that, as in the case with instrumental variables, the measure of tri-monthly inflation is the only one with a positive coefficient.

This suggests that inflation only increases as a result of the increase in the likelihood of fiscal constraint three months after the fact. In this case, it does so at a rate of 60 percentage points for every 10 percentage point increase in the likelihood of fiscal constraint. Again, this is consistent with the finding that most products change their price only every 100 days.

Price Dispersion

I do not conduct heteroscedasticity analysis of price dispersion because this would require that I calculate the variance of a variance. The kurtosis, to my knowledge, is as yet an unexplored measure of inflation; as such, I would not know how to interpret the results.

5.2.2 Expected Inflation

Table 6 presents the results for the effect of a change in the likelihood of fiscal constraint on measures of expected inflation. Recall that I utilize yields for Bocon 2016 and Bocon 2014, and hence expect to observe a negative coefficient by my hypothesis. In contrast I expect to observe a positive coefficient for Bogar 2018 since I use closing price. I also expect a positive coefficient in the fourth column since I expect inflationary expectations to increase with the likelihood of fiscal constraint.

Table 6: Results for Heteroscedasticity Analysis of Expected Inflation

	Bocon 2016	**Bocon 2014**	**Bogar 2018**	**Professional Forecast**
ΔF_t	3.284911	-1.128492	-12.93949	-.2083157
Bootstrap Standard Error	18.34976	9.140269	81.9487	44.60934
95% Confidence Interval	(-32.6, 39.2)	(-19.04, 16.79)	(-173.5, 147.6)	(-87.64, 87.22)
Events	15	15	15	10
Observations	212	212	212	36
Bootstrap Replications	982	955	999	1000

The change in the likelihood of default is calculated over a 1-day window.

All results are insignificant to the 95% confidence level and only the bond 'Bocon 2014' reflects directional changes consistent with the hypothesis. Again, a possible explanation is that changes to the expected risk of non-payment are overriding the changes to expected inflation.

6 Discussion

The results are by and large insignificant and in some cases, the point estimate suggests the opposite to the hypothesis that fiscal constraint causes inflation. This fact might be explained by one, or both of the following: there is a problem with the empirical strategy, or there is a problem with the hypothesis.

In terms of the empirical strategy, the problem may be that the case *NML Capital v. The Republic of Argentina* is not a suitable natural experiment. However, other papers (Hébert and Schreger 2016) have successfully used this source of exogenous variation before me. It is still possible that the context is inappropriate in the sense that the paper mentioned focuses on financial variables and that the extension of their strategy to a slower moving variable such as inflation is inappropriate. However, the results in Hébert and Schreger (2016) are so robust that this is unexpected: if they are observing real changes in the economy, we should see this reflected in all economic variables.

Another important difference regarding the paper by Hébert and Schreger (2016) is that, to measure the change in the Credit Default Swap spreads, they utilize data from Markit; I was unable to use this data set for financial reasons[19]. Perhaps my use of the more unreliable data set from Bloomberg introduced noise and had some bearing on the lack of significant results (attenuation bias).

Another issue of concern in the preceding methodology is that the events may not be independent and identically distributed. All the rulings preceding the June 16th, 2014 verdict by the Supreme Court were appealable. So, it is reasonable to believe that after each ruling, and knowing that there would be subsequent rulings, agents updated their beliefs on the distribution of the shocks. This Bayesian view of updating puts into question the assumption of identical distribution of shocks of Hébert and Schreger (2016) that I follow.

The second, more interesting, possibility is that the hypothesis is wrong. On the one hand, this could suggest that firms—in this case specifically the supermarket chain COTO—are not, in fact, forward-looking rational price setters. On some level,

[19]Markit gave me a quote on the data upwards of $1,000 dollars.

this contests the assumption of rational expectations. Another possible explanation is that price-setters only respond to very salient incentives, such as what is reported in the newspapers, and that the rulings of a Second District Judge in New York simply were not salient enough to affect price-setting behavior.

6.1 Actual Inflation

Some issues with the measurement of actual inflation may also explain the lack of results. First of all, despite the fact that the dataset possessed very high-frequency price information, inflation may be a slow moving variable. In this case, it is unlikely to reflect the impact of a shock immediately. In the time that it takes to react, other stimuli may have interfered.[20]

Another drawback is that the dataset focused on one retailer: COTO. Although it is a good representation of price behavior in that it is one of the largest retailers in Argentina, it is still only one source of price variation. Further iterations of the analysis may incorporate pricing information from other retailers.

6.2 Expected Inflation

Several issues with the measurement of expected inflation could explain the lack of significant findings.

The inflation-indexed bonds have the advantage of being high-frequency measurements of market expectations. However, they are indexed to official inflation. As

[20] I address this issue in the analysis by incorporating measures of expected inflation. However, as mentioned below, these measures also had their drawbacks.

was previously mentioned, despite the positive correlation between actual and official inflation, it is possible that this was not enough correlation for the pricing of these bonds to capture the effect of fiscal constraint.

More importantly, the identification assumption may be jeopardized in the case of the inflation-indexed bonds. While the price of these bonds should increase if there is higher expected inflation, they should decrease if the debtor's risk increases. Judge Griesa's rulings had both of these effects: according to the theory, unfavorable rulings both increased the expectation of inflation and increased the riskiness of the debtor (Hébert and Schreger (2016) focus on the rulings' effect on the credit-worthiness of Argentina). Hence, the rulings should both decrease and increase the yield. Ultimately, it is possible that these opposing effects are offsetting in the data and lead to insignificant coefficients.

The professional forecast, while valuable because it is created by professionals, is also an imperfect measure of the expectation of inflation for this analysis. The frequency is low, only monthly, and the variable that is forecast, annual inflation, contemplates all the events that occur throughout the year. For example, weekly forecasts of inflation for the following twelve weeks might be a cleaner measure of the impact of fiscal constraint on inflation.

7 Conclusion

The research presented in this thesis sought to contribute to our understanding of the determinants of inflation by studying the effect of fiscal constraint in Argentina during 2012-2014. In order to identify a causal relationship, I exploit exogenous variation to the likelihood of fiscal constraint, as measured by CDS spreads, arising from rulings in the court case *NML Capital v. The Republic of Argentina*. Ultimately, the evidence does not allow me to reject the null hypothesis that there is no effect of fiscal constraint on inflation.

Further research may address the issues raised in Section 6. If these issues can be resolved, then it may be worth considering the channel through which a fiscal constraint as defined here—namely a reduction in the government's ability to finance a deficit through issuing external debt—affects the domestic price level. Given the heavy reliance on the dollar in many developing economies, it is of policy interest to investigate whether a fiscal constraint affects prices through its impact on the price of the dollar, more broadly by impacting expectations of further reliance on seigniorage, or a mixture of both.

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9 Appendix A: Construction of the Inflation Index

To construct an aggregate inflation index for a discrete basket of goods, I extend the aggregate price index for a continuum of goods derived by Maurice Obstfeld and Kenneth Rogoff in Chapter 4, Section 5 of the textbook ‘Foundations of International Macroeconomics’.

The construction of this aggregate inflation index assumes Cobb-Douglas preferences. This may not be realistic since consumers do not normally substitute between strawberries and cars at the same rate as between strawberries and raspberries. However, the assumption of unitary elasticity of substitution is widely used in the literature and appropriate for this analysis.

Additionally, contrary to standard constructions of inflation indices, this index attaches equal weights to all products in the basket. This is because I am interested in studying the general increase in the price level and not how households are affected by it. If one were interested in conducting a welfare analysis as well, then the latter type of index would be more appropriate. However, this is beyond the scope of the current work.

Derivation of the Inflation Index

According to Obstfeld and Rogoff, the aggregate price index for a continuum of goods at time t is given by:

$$P_t = \exp[\int_0^1 \log p(z) dz] \tag{31}$$

Where the continuum of goods is indexed by $z \in [0, 1]$. See Obstfeld and Rogoff (1996) for a full derivation of this price index.

Hence, if aggregate inflation is given by

$$\Pi_t = \frac{P_t}{P_{t-1}}$$

Then,

$$\log(\Pi_t) = \log(P_t) - \log(P_{t-1})$$

By equation (31),

$$\log(\Pi_t) = \log(\exp[\int_0^1 \log(p_t(z))dz]) - \log(\exp[\int_0^1 \log(p_{t-1}(z))dz])$$
$$\log(\Pi_t) = \int_0^1 \log(p_t(z))dz - \int_0^1 \log(p_{t-1}(z))dz$$
$$\log(\Pi_t) = \int_0^1 \log(\frac{p_t(z)}{p_{t-1}(z)})dz$$
$$\log(\Pi_t) = \int_0^1 \log(\pi_t(z))dz$$

Hence,

$$\Pi_t = \exp[\log(\Pi_t)]$$
$$\Pi_t = \exp[\int_0^1 \log(\pi_t(z))dz]$$

The analogue for the case of a discrete basket of $i \in [1, n]$ goods is given by:

$$\Pi_t = \exp[\frac{\sum_{i=1}^n \log(\pi_t^i)}{n}] \tag{32}$$

Where inflation of good i is given by $\pi_t^i = \frac{p_t^i}{p_{t-1}^i}$.

Equation (32) is the inflation index used in the empirical analysis throughout this thesis. Note that the index is adjusted accordingly for each window of inflation. For weekly inflation for example, π_t^i becomes $\pi_t^i = \frac{p_t^i}{p_{t-7}^i}$.

10 Appendix B: The Events

Table 7 shows each ruling with its corresponding date, change in the likelihood of default over a 1-day window, daily inflation following the event and the inflation that occurred over the three month period following that event. The gaps in the data on tri-monthly inflation reflects the loss of observations due to the elimination of overlapping variables in order to avoid serial correlation. Furthermore, the red/blue

Table 7: The Events

Events	Description	ΔF (%)	Daily ∏	Tri-Monthly ∏
11/27/12	Judge Griesa denies exchange bondholders request for a stay.	2.18	1.0003585	-
11/29/12	Appeals court grants emergency stay.	-8.83	1.0008587	-
12/5/12	Appeals court denies holdout request.	-6.64	1.0005609	-
12/7/12	bondholders to appear as interested parties.	-3.36	1.0007688	1.036944
1/11/13	Appeals court denies exchange bondholders right to appeal.	3.79	1.0002139	-
3/4/13	Appeals court asked Argentina for a payment formula.	-0.31	1.0002184	-
3/27/13	Appeals court denies Argentina's request for en-banc rehearing of appeal	1.97	1	1.0325317
8/26/13	Appeals court upholds Judge Griesa's decision.	3.43	1.0007614	1.0528601
10/4/13	Judge Griesa prohibits debt swap of exchange bonds to Argentine law bonds.	0.47	1.0001175	-
10/8/13	Supreme court denies Argentina's first petition.	-0.59	1.0000856	-
11/19/13	Appeals Court denies argentina's request for an en-banc hearing.	-0.51	1.0002584	-
1/13/14	Supreme court agrees to hear Argentina's case.	1.41	1.0004133	1.1202847
6/16/14	Supreme court denies Argentina's second petition.	3.73	1.0002978	-
6/24/14	Judge Griesa prohibits debt swap of exchange bonds to Argentine law bonds.	0.88	1.001367	-
6/27/14	Judge Griesa rejects Argentina's application for a stay.	3.33	1.0004803	1.076013

coloring corresponds to unfavorable/favorable rulings for Argentina. The criteria for classification is the nature of the rulings themselves, not the observed change in the likelihood of fiscal constraint.

11 Appendix C: Further Results

This appendix presents the results for the same analysis where the change in the likelihood of fiscal constraint is calculated over a 2-day window. The reasoning for including this analysis is that Hébert and Schreger (2016) conduct their analysis using 2-day windows in their calculations. However, while these results are more readily comparable with the aforementioned authors' in terms of the methodology, they are still insignificant.

Actual Inflation

Relative Prices

Table 8: Results for Instrumental Variables Analysis Actual Inflation II

VARIABLES	(1) Daily Inflation	(2) Weekly Inflation	(3) Monthly Inflation	(4) Tri-Monthly Inflation
ΔF_t	-0.00364***	0.000237	-0.115	0.209
	(0.000388)	(0.0158)	(0.151)	(0.316)
Constant	1.001***	1.005***	1.021***	1.065***
	(0.000101)	(0.000638)	(0.00279)	(0.00911)
Observations	642	116	21	9
IV F-Test	564.39	2937.07	0.08	1.06

The change in the likelihood of fiscal constraint is calculated over a 2-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address serial correlation in the data (Newey West, 1987). However, for weak instruments, i.e. (3) Monthly Inflation and (4) Tri-Monthly Inflation, Anderson Rubin confidence intervals are more appropriate (Anderson Rubin, 1949). In this case the A-R confidence interval for both is the entire sample space.

Table 9: Results for Heteroscedasticity Analysis of Actual Inflation II

	Daily Inflation	**Weekly Inflation**	**Monthly Inflation**	**Tri-Monthly Inflation**
ΔF_t	-.0030947	-.0033254	-.455124	10.04185
Bootstrap Standard Error	.0207858	12.96017	91.06797	87.35966
95% Confidence Interval	(-.044, .0378)	(-25.4, 25.4)	(-178.9, 178)	(-161.2, 181.3)
Events	15	11	8	5
Observations	214	117	21	9
Bootstrap Replications	1000	1000	999	954

The change in the likelihood of fiscal constraint is calculated over a 2-day window. All estimations are insignificant at the 95% confidence level.

Price Dispersion

Table 10: Results for Instrumental Variables Analysis of the Variance in Actual Inflation II

VARIABLES	(1) Var(Daily π)	(2) Var(Weekly π)	(3) Var(Monthly π)	(4) Var(Tri-Monthly π)
ΔF_t	-0.000914	-0.988	-0.0597	0.507**
	(0.000780)	(6.456)	(0.139)	(0.210)
Constant	0.000588***	0.757	0.0170***	0.0485***
	(9.09e-05)	(0.705)	(0.00229)	(0.00643)
Observations	211	116	21	9
IV F-Test	201.12	2937.07	0.08	1.06

The change in the likelihood of fiscal constraint is calculated over a 2-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address serial correlation in the data (Newey West, 1987). However, for weak instruments, i.e. (3) Var(Monthly π) and (4) Var(Tri-Monthly π), Anderson Rubin confidence intervals are more appropriate (Anderson Rubin, 1949). In this case the A-R confidence interval for both is the entire sample space.

Table 11: Results for Instrumental Variables Analysis of the Change in the Cross-sectional Variance of Prices II

VARIABLES	(1) Δ_1Var(P_t)	(2) Δ_7Var(P_t)	(3) Δ_{30}Var(P_t)	(4) Δ_{90}Var(P_t)
ΔF_t	0.00547	0.149	-0.154	-0.585
	(0.00593)	(0.105)	(0.525)	(0.817)
Constant	1.001***	1.007***	1.035***	1.104***
	(0.000278)	(0.00166)	(0.00812)	(0.0175)
Observations	211	116	21	9
IV F-Test	201.12	2937.07	0.08	1.06

The change in the likelihood of fiscal constraint is calculated over a 2-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address serial correlation in the data (Newey West, 1987). However, for weak instruments, i.e. (3) $\Delta_{30}Var(P_t)$ and (4) $\Delta_{90}Var(P_t)$, Anderson Rubin confidence intervals are more appropriate (Anderson Rubin, 1949). In this case the A-R confidence interval for both is the entire sample space.

Expected Inflation

Instrumental Variables

Table 12: Results for Instrumental Variables Analysis of Expected Inflation II

VARIABLES	(1) Bocon 2016	(2) Bocon 2014	(3) Bogar 2018	(4) Professional Forecast
ΔF	9.624	20.12***	-13.54***	-0.865
	(6.333)	(1.659)	(1.011)	(2.209)
Month Dummy				-0.0572
				(0.0669)
Constant	-0.0340	0.0271	0.191***	0.561
	(0.0303)	(0.0617)	(0.0374)	(0.586)
Observations	365	286	552	30
IV F-Test	2097.57	10365.65	13552.68	32.37

The changes in the likelihood of fiscal constraint, the changes in the yield of Bocon 2016/Bocon 2014 and the changes in the price of Bogar 2018 are calculated over a 2-day window. Standard Errors are in parenthesis. ***p<0.001, **p<0.05, *p<0.1. The table reports Newey-West standard errors to address serial correlation in the data (Newey West, 1987).

Heteroscedasticity Analysis

Table 13: Results for Heteroscedasticity Analysis of Expected Inflation II

	Bocon 2016	**Bocon 2014**	**Bogar 2018**	**Professional Forecast**
ΔF_t	15.31774	20.42642	-9.159762	-1.556205
Bootstrap Standard Error	30.93547	62.11864	180.1562	45.01528
95% Confidence Interval	(-45.3, 75.95)	(-101.32, 142.17)	(-363.03, 343.17)	(-89.78, 86.67)
Events	15	15	15	10
Observations	212	212	212	36
Bootstrap Replications	982	991	999	1000

The change in the likelihood of default is calculated over a 2-day window. The change in the yields/prices of the inflation-indexed bonds matches this 2-day window