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Is Credit Dollarization Contagious across Countries? Evidence from Transition Economies

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Evidence from Transition Economies

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Abstract

We examine whether credit dollarization is contagious across countries with widely different experiences of the phenomenon and, if so, what factors contribute to such spillover effects. We analyse a unique monthly data set of credit dollarization for 23 transition economies. We proceed in two steps. First, we use a flexible bivariate regime-switching model to simultaneously test for shift contagion and bi-directional pure contagion between high-dollarized and low-dollarized countries. We document widespread evidence of both shift and pure contagion in credit dollarization, the latter moving in both directions. Second, a multivariate analysis identifies the factors that promote pure contagion to be associated with (i) geographical proximity between countries, (ii) a common institutional environment within the EU, (iii) greater trade and banking connectivity, and (iv) the economic size of the country where contagion originates from.

JEL Classification : C23; C32; E44; F31; F42; G15; G20

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1. Introduction

Foreign currency loans are among the most important features of recent financial developments in transition economies, particularly those located in Central and Eastern Europe.[1] Although there are many underlying reasons of why borrowers and creditors prefer holding foreign currency denominated assets, some are related to achieving higher stability in the value of the contracts between economic agents, especially during periods of crises.[2] This view suggests that a country's composition of loans between domestic and foreign currencies is an outcome of optimal decision-making by rational agents, so that it should not be wholly viewed as a negative development. However, the decisions of agents at the microeconomic level albeit optimal, can create concerns at the macroeconomic level when countries have a high share of liabilities in foreign currency. Adverse implications may include the heightened vulnerability of the financial sector, the effectiveness of monetary policy, and balance sheet effects (see De Nicoló et al., 2005; and Levy Yeyati, 2006). Understanding the causes of credit dollarization is important, not only for forecasting purposes, but also for designing and implementing policies to contain its size when it surpasses its desired level. Yet, with only one exception, the existing literature on the drivers of credit dollarization focuses on country-specific factors, and does not consider cross-country contagion aspects. In this paper, we study the interdependence of this phenomenon across countries and examine the channels of its transmission.

The literature that studies the determinants of dollarization has traditionally focused on its levels, which reflect a measure of its stock value. Therein, the main contributing factors have been found to be the rate of inflation in line with the currency substitution view (Savastano, 1996), the minimum variance portfolio (MVP) dollarization share according to the portfolio view (Ize and Levy-Yeyati, 2003), and the low quality of institutions according to the institutional view (De Nicoló et al., 2005). Other studies have also underscored the roles of the exchange rate regime (Arteta, 2005), interest rate differentials and access to foreign funds (Basso et al., 2011), EU membership and

[1] The European Bank for Reconstruction and Development in a recent report (EBRD, 2010) concludes that a relatively high share of foreign currency lending has been a feature of domestic bank systems in these economies for a long time.

[2] For this, see Ize and Levy Yeyati (2003) who present a portfolio model of financial intermediation in which currency choice is determined by hedging decisions on both sides of a bank's balance sheet.

expectation of euro adoption (Neanidis, 2010), and remittances (Firdmuc et al., 2013). More recently, some studies opted to investigate the causal factors of short-run variations in dollarization (Neanidis and Savva, 2009; Basso et al., 2011). Their rationale is that the stock measures of dollarization are very persistent, so that they do not capture well the changes in dollarization due to new business activity. Flow measures instead explain the portfolio choice of agents on new loans and deposits, thereby arguing that agents do not re-balance their entire portfolios at every period but perhaps just renew loans and deposits previously done. This period-by-period consideration of the agents' portfolio choice has shown that short-run dollarization is particularly prone to exchange rate changes, changes in monetary expansion, and to the banks' currency matching behavior of loans and deposits.

Whilst the foregoing research has yielded valuable insights on the main drivers of dollarization, there is still considerable room for further investigation. A particularly fertile area, which so far has gone undetected, is the extent to which dollarization is subject to contagion effects beyond national borders. An attempt in this direction has been made by Neanidis and Savva (2018) in a companion study (henceforth NS), who took a step beyond the country factor-specific analysis of dollarization towards the relatively unexplored cross-country spillover analysis of financial dollarization. Employing a panel data model that uses spatial econometric techniques, NS study the regional spillovers of deposit and credit dollarization for a group of 23 countries. The countries belong in the same geographic region, Central and Eastern Europe, and the period under investigation extends over two decades, 1990s and 2000s. The first finding of the analysis is that both deposit and loan dollarization spill over across countries. On average contagion is found to raise country-specific dollarization by 3.7 percentage points in deposits and by 4.7 percentage points in loans. The second finding is that contagion is transmitted via bilateral trade linkages among countries and the degree by which the domestic banking system is linked to foreign banks, both encouraging cross-border lending.

Prompted by the findings of NS, in this paper we take the investigation a step further. Our contribution is threefold. First, we examine contagion in short-run credit dollarization given the recent interest of the literature in this measure of dollarization.

Our use of monthly data permits focusing on new business activity in dollarization in an effort to move away from its time-persistent stock measure. From an intuitive point of view, we think it makes sense to estimate contagion in short-term dollarization since a flow measure reacts more forcefully to shocks compared to its stock counterpart. In this way, the cross-country transmission of short-run dollarization is following the propagation of shocks across countries. Second, the investigation takes account of contagion effects between very different countries when it comes to their experience of dollarization. We split our countries into two groups according to their historical levels of dollarization, high-dollarized and low-dollarized.[3] In this way we can examine heterogeneous effects in cross-country dollarization spillovers and check whether one country group exerts more influence to the other. The focus on this type of heterogeneity follows from the findings of NS who show the magnitude of the dollarization spillover to be greater *within* the group of low-dollarized countries. Here we explore whether spillovers in dollarization differ in size *across* high-dollarized and low-dollarized countries. This is an important aspect of the analysis because if the group with low (high) levels of credit dollarization reacts proportionately more to contagion, this implies convergence (divergence) in credit dollarization between country groups. Third, we examine the factors that contribute to cross-country contagion effects in an effort to identify the channels of contagion. In addition to the banking and trade linkages identified by NS, predictors include geographic closeness, institutional structure, and the economic size of the country where contagion stems from.

To achieve our objectives, we proceed in two steps. In the first step, we apply a regime-switching model that tests for two distinct channels of contagion between countries in the two groups: shift contagion and pure contagion (Gravelle et al., 2006; Flavin et al., 2008; Flavin et al., 2014). The former occurs when there is a change in the normal levels of interdependence between countries during turbulent times due to common shocks. The latter takes place during crises not explained by market fundamentals and common shocks. An important qualification of our analysis is that we allow for a simultaneous test for pure contagion running in both directions, from high-

[3] Neanidis and Savva (2009) also show the effects on short-run dollarization of local-currency depreciation and monetary expansion to differ in countries with different levels of dollarization.

dollarized to low-dollarized countries and from low-dollarized to high-dollarized countries. Accounting for these bi-directional effects allows fully capturing the impact of interactions across country groups. In the second step, we explore the reasons underlying the transmission of contagion across countries. In identifying the channels of contagion, we use the estimates of contagion obtained in the first step as dependent variable in a specification with aggregate predictors.

We find strong evidence in favor of shift contagion and of bi-directional pure contagion between high-dollarized and low-dollarized countries. Analysis of the relative impact of the two types of contagion reveals that for both the common and the idiosyncratic shocks, it is the group of low-dollarized countries that react proportionately more to these shocks. This supports the argument that credit dollarization is subject to convergence "from below." Beyond the evidence of contagion, our identification strategy reveals that the magnitude of the spillover effect is dictated by the countries' proximity to each other, the sharing of a common institutional framework under the structure of the European Union, and macroeconomic indicators of international linkages and economic size. Overall, the results highlight the key role of international coordination as an instrument in assisting countries to achieve their individual desired levels of credit dollarization.

Our contribution is related to two strands of the existing literature. It is linked to the literature that examines the determinants of foreign currency lending (Luca and Petrova, 2008; Neanidis and Savva, 2009, 2013; Neanidis, 2010; Basso et al., 2011; Firdmuc et al., 2013; Kishor and Neanidis, 2015). With the exception of NS, none of these studies considers cross-border spillovers in credit dollarization. We complement NS by examining two specific types of contagion, shift and pure, that permit testing for cross-country convergence of the phenomenon. Also, our study extends the analysis of the channels of transmission to more international linkages.

Our study is broadly related to the literature that examines contagion effects in financial assets across countries. Since the early studies that tested contagion across financial markets (i.e., King and Wadhwani, 1990; Lee and Kim, 1993; Forbes and Rigobon, 2002), an explosion of recent work has documented shift and pure contagion effects in financial instruments.[4] For example, Gravelle et al. (2006) investigate shift contagion in the currency and bond markets of various developed and emerging-market countries, while Flavin et al. (2008) test for both shift and pure contagion in equity and currency markets in a group of East Asian emerging economies. In our setting the distinctive characteristic of the analysis is the type of the financial asset, where we draw attention to contagion effects of an asset particularly relevant to the banking systems of Central and Eastern Europe.

The remainder of the paper is organized as follows. Section 2 describes the data and the econometric model. Section 3 presents the main findings of the analysis with regard to the tests of contagion and to the economic identification of the contributing factors of pure contagion. Section 4 concludes.

2. Data and econometric model

2.1 Data

The objective of this study is to build on the findings of NS to examine cross-border spillovers in credit dollarization (CD) between countries that have different dollarization experiences, low vs. high. To be consistent, we use the same country sample and period coverage for CD as NS. This corresponds to 23 transition economies located in Central and Eastern Europe and the former Soviet Union since the early 1990s until the end of the 2000s. The sample includes Albania, Armenia, Belarus, Bulgaria, Croatia, Czech Republic, Estonia, Georgia, Hungary, Kazakhstan, Kyrgyz Republic, Latvia, Lithuania, Macedonia FYR, Moldova, Poland, Romania, Russia, Serbia, Slovak Republic, Slovenia, Turkey and Ukraine. The choice of the sample has been guided by two principles to best capture contagion and spillover effects: (i) the common geographic region countries belong to, and (ii) their comprehensive coverage of the CD series. Table A1 in the Appendix lists all countries with their respective period coverage and number of observations.

[4] An excellent guide on the taxonomy and the various categories of international financial contagion is provided by Moser (2003).

Consistent with the literature, CD is defined as the ratio of foreign currency credits to total credits of domestically-based banks to a country's residents. This definition highlights that this measure of dollarization includes credits in every type of foreign currency, not just dollar or euro holdings. The data are at monthly frequency and primarily drawn from National Central Bank reports, yielding a total of 3653 observations.[5] Table 1 (upper panel) provides summary statistics for CD over the entire sample and by splitting the sample between high- and low-dollarized observations with the threshold been drawn at the mean value of CD, located at 41.49%. This value is close to the median value of the sample (41.05%) and to the threshold of 40% both used in the literature to separate countries with low and high levels of dollarization (Levy-Yeyati, 2006; Rennhack and Nozaki, 2006; Neanidis and Savva, 2009). In our sample, the mean value yields two sub-samples of almost equal size: 1810 vs. 1843 observations.

Notable characteristics related to the level of CD are as follows. First, the mean value is relatively high with substantial variation across countries. Second, the sample mean does a good job separating the two sub-samples of countries in terms of the size and the dynamics of CD. Both these characteristics are reflected in Figures 1a and 1b which present the patterns of CD across high-CD (HCD) and low-CD (LCD) countries. The former group includes 11 countries and the latter 12 countries. The mean CD value, shown with a solid horizontal line in both figures, draws a clear demarcation between the two country groups. Importantly, countries in each group appear to stay in their own group and to rarely move to the other group. In other words, CD exhibits high persistence across time. This country grouping, we claim, can allow us testing for the presence of contagion in CD across HCD and LCD countries.[6]

[5] The countries in our sample differ in economic size and this could partly explain their different levels of CD. Following the literature, we use in our analysis the actual dollarization ratios rather than construct weighted values by economic size. We opt including the latter as control in a regression that explains the variation in pure contagion across countries, see Section 2.2.2 below.

[6] Table A1 indicates that the countries in our sample have different number of observations due to varying time period coverage. This means that in each country group there are countries with shorter time series than others. For this reason, one cannot treat all countries in a group as exerting the same weight to the group because this can bias the measurement of CD within each group due to missing observations. To avoid this problem in the analysis, similar to NS, we use the level of CD calculated as the weighted average of the observable data, where the weight of each country reflects its importance to the total CD of the group it belongs to.

Since our main goal is to examine heterogeneous effects in cross-border contagion, we divide the sample into two groups of countries with different levels of CD. However, pooling the data into two groups does not come without a cost. This corresponds to the loss of cross-sectional variation that can help identify cross-country contagion. Moreover, some of the countries that are being grouped together are so different to each other, that the mechanical sorting and aggregating into two groups cannot easily yield intuition as to why CD spills over from one country group to the other. For this reason, we also apply a strategy where we disaggregate the two groups and repeat the analysis at the country level for pairs of countries. Each pair contains a country from each group, giving rise to 132 (=11*12) country pairs. This yields a richer set of estimates that can provide more insights about the degree of contagion across countries.[7] Importantly, we make an effort to identify a causal relationship of cross-country contagion by grounding the transmission of CD to economic, institutional and proximity linkages between pairs of countries (more on this in the next section).

Figures 1a and 1b suggest a further piece of information: the CD series is not stationary for any of the 23 countries. As the models that test contagion effects at the time-series level are based on stationary variables, given that the use of nonstationary variables may create inference problems (Gravelle et al. 2006; Flavin et al. 2014), we formally test for stationarity by performing the Ng–Perron unit root test for each country in the sample. The test suggests that we cannot reject the null of a unit root at conventional significance levels for any country, concluding that the level of CD is nonstationary.[8] For this reason, we use the first differences of the CD series, which are found to be stationary. This measure is defined *short-run* credit dollarization, a term used by Honohan (2007) and Neanidis and Savva (2009). The lower panel of Table 1 presents

[7] In these bivariate combinations, the estimations use only the matched data sample of each pair of countries under study. That is, they use the common time series period available for the two countries in a pair. In this way, there is no need for weighing the series as done in the analysis that uses groups of countries.

[8] We obtain the Mza statistic of the test, with the 5% critical value determined at -8.1 and the number of lags determined by the SIC criterion. Other unit root tests, such as Phillips–Perron and Elliott–Rothenberg–Stock, also do not reject the null of a unit root at any level of significance. Notably, this finding still holds when using the critical values of Cavaliere (2005) who corrects for time series that are bounded from both below and above, as our CD series is. All results are available upon request.

summary statistics associated with this variable and Figure 1c offers a graphical illustration.

The first difference in CD (ΔCD) shows that, on average, CD increases by 0.1% across the sample. This figure jumps to 0.17% in the HCD sub-sample and is only 0.02% in the LCD group. This further supports the claim of two distinct country groupings in CD that grow at different rates. Unsurprisingly, the volatility of short-term CD is higher in the HCD group. To get an economic feeling for the magnitude of the average short-term CD, we consider a representative country and calculate the level of its CD in the beginning and at the end of the sample period. We choose as the representative country, the "mean country" in the sample, i.e., one with mean CD of 41.49% and mean ΔCD of 0.1%, the latter giving rise to monthly CD growth rate of 0.32% (or annualized growth rate of 3.9%). The start of the sample period is January 1993 and the end is November 2009, consistent with the max period coverage in our sample, covering 203 months. The monthly CD growth rate over these 203 months corresponds to a growth rate of 91.28% (calculated as $1.0032^{203} - 1$). This implies that the representative country experiencing a change in CD of 0.1%, with a CD ratio of 28.5% in January 1993, by the end of 2009 it records a CD ratio of 54.5%. This is indicative that a monthly increase of just 0.1% in CD over a 17-year period, results into an almost doubling of the original magnitude in CD. Effectively, this makes the representative country switch from a low-dollarized economy into a high-dollarized one.

Completing the description of Table 1, statistics on the short-run CD series include results on skewness, kurtosis and Jarque-Bera normality tests which provide evidence against normality in ΔCD across countries. In addition to the above, we perform diagnostic tests on standardized residuals. The Ljung–Box test for up to twelve lags serial correlation indicates the strong presence of serial dependence in the data. Similarly, a Ljung–Box test for serial correlation in the squared data provides evidence of conditional heteroskedasticity.[9] All these tests suggest that short-run CD is best modelled as a mixture of distributions, supporting the use of a model specification that allows for more than one volatility regime, like the one presented in the next section.

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[9] To save space all diagnostic tests are available from the authors upon request.

2.2 Econometric model

The analysis has two objectives. First, test for the presence of shift and pure contagion in CD across (groups of) countries characterized by different levels of CD. Second, offer a causal interpretation of contagion by identifying its underlying determinants. To achieve these objectives we resort to the Estimated Dependent Variable (EDV) approach first described by Hanushek (1974) and more recently by Lewis and Linzer (2005). This involves two steps, corresponding to each of the two objectives. In the first step, we estimate a factor model for two groups (or pairs) of countries to obtain estimates of cross-border contagion. In the second step, the estimates of contagion obtained in the first step become the dependent variable explained by a set of aggregate predictors. The EDV approach is ideal for our analysis because it allows estimating complex models in the first step that are difficult to estimate using multilevel techniques. In addition, in our case the first step involves a time series analysis based on monthly data, whereas the second step makes use of the cross-sectional variation amongst country pairs. Due to these, there is no technique that allows incorporating the two steps into one. Below we describe each step in detail.

2.2.1 Testing for shift and pure contagion: a bivariate analysis

The objective of the first step is to test for the presence of both shift and pure contagion in CD between two groups or pairs of countries: HCD countries and LCD countries. To do this, we employ a methodology developed by Gravelle et al. (2006) that captures shift contagion and its extension by Flavin et al. (2008) that accounts for pure contagion and spillover effects in international asset markets (e.g., bonds, currencies, equity, gold). This framework is well-suited to capturing the different channels of contagion and has been used widely in the financial and macroeconomic analyses of contagion (e.g., see Grier et al. 2004; Bredin and Fountas 2009; Neanidis and Savva 2013; Flavin et al. 2014). The model belongs to the family of factor models and is bivariate in nature.[10]

[10] We choose a bivariate framework compared to a multivariate analysis because the latter would lead to an explosion in the number of parameters that need to be estimated, thus, making its estimation impossible. Nevertheless, we try to get the most out of both frameworks by complementing the bivariate analysis of the first step with a multivariate analysis in the second step.

The model in the first step uses dynamic factor analysis, which dictates the use of stationary variables. Given this requirement, we opt using in the analysis the first difference in CD as our variable of interest. Under this framework, the ΔCD in country group *i* at time period *t* can be decomposed into two components as follows

$$\Delta CD_{it} = \mu_{it} + u_{it}, \tag{1}$$

where μ_{it} is the expected change in CD in the two groups of HCD and LCD countries, $i = H, L$, and u_{it} is the unexpected component, with $E(u_{it}) = 0$ and $E(u_{Ht}, u_{Lt}) \neq 0$. The separate treatment of HCD and LCD countries in the model is supported by the different rate by which CD changes across the two country groups, as documented in the previous section. The analysis applies equally to pairs of countries by pairing a country from the HCD group with a country from the LCD group. Equation (1) further assumes that the changes in CD are driven by two factors, one predictable, μ, and one unpredictable, u.

Because the two components are not observable, we need to specify a dynamic structure for their identification. For this reason, we specify two processes as follows. Starting with the unexpected components, u_{it}, it is assumed that there exists contemporaneous correlation between them. This, in turn, implies that common structural shocks may be driving both country groups' short-run CD behavior. As such, u_{it} is decomposed into two structural shocks, a common (z_{ct}) and an idiosyncratic (z_{it}). The common shock captures the impact of a CD innovation that is common to both groups of HCD and LCD countries, while the idiosyncratic shock is unique to each individual country group. The process is described as

$$u_{it} = \sigma_{cit} z_{ct} + \sigma_{it} z_{it}, \tag{2}$$

where σ_{cit} and σ_{it} measure the impact of the two structural shocks on each country groups' CD changes. To facilitate an easier interpretation of the coefficient estimates in equation (2), both shock variances are normalized to unity. In this way, σ_{cit} and σ_{it} represent standard deviations of each respective shock.

Further, since we are interested in testing for contagion in short-term CD across HCD and LCD countries, we allow both the common and the idiosyncratic shocks to switch between two states of volatility, high and low.[11] A comparison between the impact coefficients of the shocks across the two volatility regimes in the two country groups allows for a formal test of contagion. Specifically, following Gravelle et al. (2006), the regime-switching behavior of the common shock is sufficient to identify the parameters associated with shift contagion, while, according to Flavin et al. (2008), the regime-switching behavior of the idiosyncratic shock identifies the pure contagion parameters. We assume that the structural impact coefficients, σ_{cit} and σ_{it}, switch between the two volatility states according to

$$\begin{aligned} \sigma_{cit} &= \sigma_{ci}(1-S_{ct}) + \sigma_{ci}^{*}S_{ct}, \\ \sigma_{it} &= \sigma_{i}(1-S_{it}) + \sigma_{i}^{*}S_{it}, \end{aligned} \tag{3}$$

where $S_{jt} = (0,1)$, $j = i, c$ are state variables that take a value of zero during tranquil times and a value of one during unstable periods. This means that values with (without) an asterisk indicate values consistent with the high (low)-volatility state.

To complete the description of the process that drives the behavior of the unexpected components, we need to specify how the volatility regimes evolve over time. Following Hamilton (1989), we assume that the regime paths are governed by a Markov switching process, where changes in volatility are endogenously determined, as follows

$$\Pr\left[S_{j,t} = 0 \middle| S_{j,t-1} = 0\right] = q_j, \tag{4}$$

$$\Pr\left[S_{j,t} = 1 \middle| S_{j,t-1} = 1\right] = p_j. \tag{5}$$

Notably, this procedure allows for sudden jumps between high and low volatility regimes.

Moving to the description of the expected component of equation (1), μ_{it}, we assume that it is time-varying and depends on the state of the common shock, (e.g. Flavin et al., 2008 and Flavin et al., 2014) as

$$\mu_{it} = \mu_i(1-S_{ct}) + \mu_i^{*}S_{ct}. \tag{6}$$

In this way, the model suggests that part of the expected short-run CD varies with the level of volatility associated with the common shock. Notice that the expected component

[11] Note that in this way, we assume the idiosyncratic shock to be unique to each country group. Ideally, and more realistically, one should consider idiosyncratic shocks that are unique to each country. This is a task we return to later in the analysis when examining pairs of countries.

is not allowed to vary with the volatility regime of the idiosyncratic shock, since these shocks are not correlated with the common shocks. But why would the volatility of short-term CD causally increase its expected value? We claim the volatility of CD mirrors the minimum variance portfolio (MVP) dollarization ratio, developed by Ize and Levy Yeyati (2003), which increases with inflation volatility and decreases with the volatility of real exchange rate depreciation. In this way, relative macroeconomic uncertainty between inflation and foreign exchange risk captured by the MVP dollarization share proxies for the volatility of credit dollarization, which has been found by numerous studies to be a strong predictor of actual dollarization ratios (Levy Yeyati 2006).[12] An advantage of our interpretation of the volatility of CD is that data on inflation and domestic currency depreciation, and hence of their volatilities, are readily available for borrowers to track so that they can adjust their portfolio of short-run loans between domestic and foreign currencies.[13] For this reason, the volatility of short-run CD by mirroring the volatility of the inflation rate to that of the real exchange rate augments expected short-term CD.

The final ingredient in order to complete the model, and allow testing for pure contagion, requires the implementation of the channels via which the idiosyncratic shock of one country group spills over to the other country group during turbulent times, over and above the effect captured by the common shock. As we do not wish to restrict ourselves to a uni-directional spillover from either the HCD countries to the LCD countries, or vice versa, we follow Flavin et al. (2008) and implement a specification that considers bi-directional pure contagion effects. This is modeled by extending, during the high-volatility regime, equation (1) for the short-run CD of the HCD country group with the idiosyncratic shock of the LCD country group and, simultaneously, augmenting the short-run CD of the LCD country group with the idiosyncratic shock of the HCD country group. A spillover from such a shock represents instability in the system, which arises due to the new channels of transmission being active during high-volatility periods. This gives rise to a total number of eight distinct regimes.

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[12] Basso et al. (2011) have shown the MVP ratio to promote CD particularly in the short-run not only for its total value, but also when disaggregated to its individual and corporate components.

[13] To give an indication of the explanatory power of the MVP share, in our companion study we find it to explain a significant fraction of the variation in CD: a one standard deviation increase in the MVP share from its mean translates into a 6 percentage point increase in CD. Given the sample average of CD of 41.5 percent, the MVP share represents a 15 percentage point contribution to CD.

The two extreme states of the model correspond to those of (i) tranquil periods and (ii) unstable periods. The tranquil periods, where *all* shocks are in the low-volatility regime, yield the following equations for short-term CD for each country group

$$\Delta CD_{Ht} = \mu_H + \sigma_{cH} z_{ct} + \sigma_H z_{Ht},$$
$$\Delta CD_{Lt} = \mu_L + \sigma_{cL} z_{ct} + \sigma_L z_{Lt}, \tag{7}$$

where the idiosyncratic shocks are assumed to be independent. In this way, co-movements in short-term CD are solely determined by the common shock and the covariance matrix of short-term CD is given by

$$\sum\nolimits_1 = \begin{bmatrix} \sigma_H^2 + \sigma_{cH}^2 & \sigma_{cH}\sigma_{cL} \\ \sigma_{cH}\sigma_{cL} & \sigma_L^2 + \sigma_{cL}^2 \end{bmatrix}.$$

During unstable periods, on the other hand, where *all* shocks are in the high-volatility regime, the model reduces to

$$\Delta CD_{Ht} = \mu_H^* + \sigma_{cH}^* z_{ct} + \sigma_H^* z_{Ht} + \delta_H \sigma_L^* z_{Lt},$$
$$\Delta CD_{Lt} = \mu_L^* + \sigma_{cL}^* z_{ct} + \sigma_L^* z_{Lt} + \delta_L \sigma_H^* z_{Ht}, \tag{8}$$

where now the variance-covariance matrix of short-run CD is given by

$$\sum\nolimits_8 = \begin{bmatrix} \sigma_H^{*\,2} + \sigma_{cH}^{*\,2} + \delta_H^2 \sigma_L^{*2} & \sigma_{cH}^* \sigma_{cL}^* + \delta_L \sigma_H^{*\,2} + \delta_H \sigma_L^{*2} \\ \sigma_{cH}^* \sigma_{cL}^* + \delta_L \sigma_H^{*\,2} + \delta_H \sigma_L^{*2} & \sigma_L^{*2} + \sigma_{cL}^{*\,2} + \delta_L^2 \sigma_H^{*\,2} \end{bmatrix}.$$

A direct comparison of equations (7) and (8) reveals that the coefficients δ_H and δ_L capture the magnitude and significance of spillovers during episodes of high volatility in the idiosyncratic shocks arising in the two country groups. Further, it is important to note that although equations (7) and (8) report, for simplicity, two extreme states as potential channels of instability, the model described by equations (1)-(8) is estimated as a complete system with eight distinct regimes. Finally, to estimate the model through maximum likelihood, we need to make an assumption that the structural shocks are normally distributed.[14]

Having described our model, we can now discuss the tests we conduct to investigate the presence of shift and pure contagion in CD between HCD countries and

[14] The model estimates in total 14 parameters whilst the country with the smallest sample in the dataset consists of 73 observations (Serbia). Despite the relatively small sample size for this country, even in this case there exist sufficient degrees of freedom to run the procedure and for the system to converge and provide meaningful estimates.

LCD countries. We test for *shift* contagion by examining the stability of the structural transmission of the common shocks as this is reflected by the impact coefficients on the common shocks, σ_{cit} and σ_{it} (Flavin et al., 2014). It is expected that during crises the common shocks, captured through countries linkages, will be larger. But, whether these increases represent a change in the structural transmission mechanism of common shocks (i.e., a change in the interdependence between the two country groups), is determined by the relative size of the common shocks between country groups. In other words, this depends on the ratio σ_{cH}/σ_{cL} . If this ratio is different between normal and turbulent times, this is an indication of shift-contagion. If, on the other hand, the ratio remains unchanged, no shift-contagion occurs. The presence of shift-contagion can be formally tested using the likelihood ratio statistic for the following hypotheses

$$H_0 : \frac{\sigma^*_{cH}}{\sigma^*_{cL}} = \frac{\sigma_{cH}}{\sigma_{cL}} \text{ vs } H_1 : \frac{\sigma^*_{cH}}{\sigma^*_{cL}} \neq \frac{\sigma_{cH}}{\sigma_{cL}},$$

where the (null) alternative hypothesis corresponds to (no) shift-contagion.[15] Essentially, the test investigates whether the higher variance and co-movement of short-run CD between HCD and LCD countries during turbulent times are due to increased impulses stemming from the common shocks, or from changes in their propagation mechanism.

Under the null hypothesis, the impact coefficients in both the tranquil and turbulent periods move proportionately so that their ratio remains unchanged. By computing a statistic γ

$$\gamma = \frac{\sigma^*_{cH}\sigma_{cL}}{\sigma^*_{cL}\sigma_{cH}},$$

we conclude that the transmission of the common shock is stable when $\gamma = 1$, which corresponds to a change in the size of the shocks only, while shift contagion exists when $\gamma \neq 1$. In the latter case, when $\gamma < 1$ there is a stronger reaction to common news from the LCD group, while $\gamma > 1$ implies a stronger reaction to common shocks from the HCD group.[16]

[15] The likelihood ratio statistic has a Chi-square(1) distribution under the null hypothesis of no shift-contagion.

[16] During the period we investigate, a small number of countries in our sample have opted for a fixed exchange rate regime. Flavin et al. (2008), however, argues that fixed exchange rates bias the results towards finding shift contagion. Relevant to the current study, Neanidis and Savva (2009) show that

We examine for *pure* contagion across country groups by testing for the stability in the structural transmission of the idiosyncratic shocks from one group of countries to the other. For this purpose, we perform a simple *t*-test on the δ_i coefficients, where under the null there is no pure contagion, $\delta_i = 0$. When this value is different from zero, however, a positive value of the parameter δ_H implies that a negative high-volatility shock, in the form of higher ΔCD, arising in the group of LCD countries, depresses short-run CD in the group of HCD countries. A negative value of δ_H, on the other hand, implies that a higher short-run CD shock in the group of LCD countries, increases short-term CD in the group of HCD countries. Analogous interpretations correspond to the parameter δ_L. That is, a negative (positive) value of δ_L implies that a higher short-term CD shock in the group of HCD countries, increases (decreases) short-run CD in the group of LCD countries. According to this interpretation, pure contagion is observed when $\delta_i < 0$, while a $\delta_i > 0$ indicates a substitution away from foreign-currency denominated loans towards loans denominated in domestic currency. In the case where both δ_H and δ_L are significant and of the same sign, we conduct a likelihood ratio test of their relative magnitude to assess which country group exerts a more sizeable spillover effect to the other.

The first step of the analysis described above, although very useful in identifying the presence of spillovers across countries, it may not reflect a causal relationship from one country group to another but merely some heightened correlation in turbulent times. It is therefore important to address identification and establish the linkages that allow shocks in CD to cross-over to other countries. To achieve this, we need a larger number of estimates of contagion that go beyond country groupings. To this extent, we repeat the above analysis by pairing every country from the HCD group with every country in the LCD group. This gives rise to 132 country pairs, each with its own estimate of contagion

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exchange rate intervention by authorities toward less flexible exchange rate regimes have no discernible effect on short-run CD. They also find that although depreciation induces banks to raise CD in the short-term, this behavior does not differ between countries with high and low levels of CD. We view these findings as evidence that fixed exchange rates do not play a role in determining short-run CD, especially for countries with different experiences in dollarization, so that a greater reaction to common shocks from any country group cannot be due to the type of the implemented exchange rate regime therein. Hence, although we acknowledge the concerns of Flavin et al. (2008), we do not view them as daunting in our framework.

that we use in a second step as dependent variable to explore the channels via which CD shocks transmit across countries.

2.2.2 The economic identification of pure contagion: a multivariate analysis

The second step of the analysis involves identifying the predictors of contagion so as to establish a causal interpretation in the transmission of CD. This amounts to using the coefficient estimates of pure contagion from the first step as dependent variable in a regression specification that considers its explanatory factors.[17] For the second step to be meaningful, we resort to the pairwise estimates of contagion because the use of country groupings produces only one such coefficient estimate. The model in the second step takes the form:

$$\begin{aligned}\delta_{H,jk} &= \alpha + \beta_1 Contiguity_{jk} + \beta_2 Distance_{jk} + \lambda_1 EU_j + \lambda_2 EU_k + \lambda_3 EU_{jk} \\ &\quad + \varphi_1 BilTrade_{jk} + \varphi_2 ForBankPenet_k + \varphi_3 logGDP_k + \varepsilon_i, \\ \delta_{L,jk} &= \alpha + \beta_1 Contiguity_{jk} + \beta_2 Distance_{jk} + \lambda_1 EU_j + \lambda_2 EU_k + \lambda_3 EU_{jk} \\ &\quad + \varphi_1 BilTrade_{jk} + \varphi_2 ForBankPenet_j + \varphi_3 logGDP_j + \varepsilon_i,\end{aligned} \tag{9}$$

where the dependent variables $\delta_{H,jk}$ and $\delta_{L,jk}$ are the estimates of pure contagion from equation (8) between pairs of countries *jk* with *j* denoting a HCD country and *k* a LCD country.

The control variables, the construction of which appears in Table A2, correspond to three types. First, following the spatial econometrics literature that examines channels of cross-sectional interdependence, we use measures of closeness between country pairs. These are generally related to geography and space and more specifically proxied by contiguity and distance, with contiguity expected to raise contagion while distance to reduce it. Given that pure contagion is captured by $\delta_i < 0$, an increasing effect of contiguity on contagion is depicted by a negative coefficient estimate of β_1. In the same manner, greater distance between countries that reduces the likelihood of contagion is shown by a positive coefficient estimate of β_2.

[17] The second step focuses on the determinants of pure contagion rather than those of shift contagion. We think this question is more interesting since pure contagion represents the transmission of country-specific shocks, the drivers of which can be proxied best by using bilateral links between country pairs. In addition, pure contagion is closer to the definition of cross-country spillovers used in the analysis of NS to which we wish to remain close for ease of comparison.

The second type of control variables are institutional in nature, proxied by dummies associated with a country's European Union (EU) membership process. When joining the EU, the expectation is that the institutional framework instilled upon EU members via their (lengthy) admission procedure, homogenizes the regulatory environment and operations in various sectors of an economy, including the banking sector. The resulting reduction in transaction costs and greater opportunities for cross-country economic interactions within the Union, leads to greater economic synchronization (see, for instance, Savva et al. 2010). Specific to spillovers in CD, Neanidis (2010) shows that EU membership by leading eventually to euro adoption, spurs a higher degree of lending in foreign currency due to the expectation of diminishing currency risk and full access to foreign currency holdings as the prospective EU members lift their restrictions on capital mobility. Through this channel, EU membership is expected to raise the degree of CD contagion. Our sample includes both EU members and non-EU countries that allow testing for this effect by adding separate dummies for each country's EU membership (EU_j and EU_k) and a joint dummy when both countries in a pair are EU members (EU_{jk}).

Finally, the third type of control variables in equation (9) includes macroeconomic indicators through which contagion can occur. These are the two instruments of cross-country spillovers in CD used by NS: (i) bilateral trade between country pairs ($BilTrade_{jk}$) and (ii) a country's share of foreign bank penetration ($ForBankPenet_k$ or $ForBankPenet_j$). To this we add a country's economic size, measured by the logarithm of its GDP. Bilateral trade is meant to capture trade linkages that reflect aggregate demand effects, while foreign bank penetration represents financial linkages that reflect credit supply channels (Forbes, 2012). Both bilateral trade and the share of domestic banks with assets of foreign ownership in excess of 50% are expected to contribute to higher levels of cross-country contagion in CD. The former by offering to countries opportunities to import and export goods and services in exchange for (foreign) currencies that induces traders in both countries to obtain loans in foreign currency. The latter by raising the exposure of a country's banking system to the international financial network, contributing to a greater transmission of shocks from one country to the next, including changes in CD (Cocozza and Piselli, 2010; De Haas and van Horen, 2012).[18] The size of GDP is also included to proxy for the "gravity" imposed by that nation between a pair of countries where contagion originates from, expected to exert a greater impact the larger its size is (similar in reasoning to the gravity equation model in international trade).[19] Following up from the above description of macroeconomic controls, all three variables are expected to enhance pure contagion, so that the estimates of φ's all take up negative values.

The cross-sectional model described in equation (9) can be estimated by OLS. However, if the sampling uncertainty in the dependent variable is not constant across observations, the regression errors will be heteroscedastic and OLS will introduce inefficiency and may produce inconsistent standard error estimates.[20] To overcome this problem and achieve significant gains in efficiency, Hanushek (1974) proposes the use of a feasible generalized least squares (FGLS) estimator. This estimator accounts for cross-sectional heteroscedasticity and corrects for bias in the standard error estimates, producing consistent estimates for the standard errors. For this reason, we estimate equation (9) by generalized least squares, incorporating the variance-covariance matrix of the parameters estimated in equation (8) (for a recent application of this procedure, see Eichholtz et al. 2010).[21]

[18] As in NS, we do not use data on bank penetration between pairs of countries because these are not readily available. Instead, we use a measure of the degree of a country's foreign bank penetration, without considering the country of origin of the penetrating bank. This effectively assumes that for any country in the sample, the transmission of CD depends on the degree of bank penetration of every other country in the region. That is, the more the region is penetrated by foreign banks, the greater the degree of transmission to each of the countries in the region. We believe this is a very plausible assumption, especially when countries share foreign ownership by international banks.

[19] Since foreign bank penetration and GDP are not bilateral variables, we include in the regression the variable that corresponds to the country where contagion emanates from. For example, when examining the determinants of δ_H which measures contagion from LCD to HCD countries, we include in the regression the values of foreign bank penetration and of GDP for the LCD countries, indicated by k.

[20] Effectively, using estimated values from a first stage as a dependent variable in a second stage introduces biased standard errors because the second step ignores the estimation error from the first stage. We cannot assume that the difference between the true value of the dependent variable and its estimated value remains the same across all country pairs, so it is very likely that the sampling variance in the δ's varies across observations. This introduces a sampling error that cannot be corrected with simple OLS.

[21] Lewis and Linzer (2005) in a series of Monte Carlo experiments provide ample evidence of the superiority of FGLS against both OLS and WLS (weighted least squares).

3. Findings

In this section, we first present the bivariate results for the two groups of HCD and LCD countries. Specifically, we report the estimates associated with the expected component of short-run CD in the two volatility regimes and with the unexpected components, the latter giving rise to the tests for shift and pure contagion. After discussing the findings with regard to shift and pure contagion at the country group level, we present the estimates of pure contagion for the 132 pairs of countries. It is these estimates then used in the multivariate cross-sectional analysis, where we establish the causal linkages of pure contagion from HCD to LCD countries and vice-versa.

3.1 Mean estimates of short-run credit dollarization

Table 2 reports the expected mean values of short-term CD across the two regimes of the common shock. Specifically, μ_H and μ_L report the expected mean values for the groups of HCD and LCD countries in the low-volatility regime, while μ^*_H and μ^*_L refer to the corresponding estimates for the high-volatility regime. The low-volatility regime is characterized by a mean short-run CD that is not statistically different from zero in the HCD countries and a negative value in the LCD countries, while the high-volatility regime indicates a positive mean short-run CD in the HCD countries and a statistical value of zero in the LCD countries. Putting these results together, they reveal that a move from the low-volatility to the high-volatility regime leads to an increase in short-run CD and, hence, raises the *level* of CD for both sets of country groups, highly dollarized or otherwise. Put differently, high volatility periods induce agents (banks and borrowers) to switch to foreign-currency denominated credits across the board.

Although we cannot directly test for the underlying factors causing this behavior, the relevant literature offers some guidance. A higher volatility is indicative of greater uncertainty and higher perceived risk for credit market participants, typically an outcome of an increase in the volatility of inflation compared to the volatility of the real exchange rate (termed MVP dollarization). A number of studies have shown that this factor contributes to higher levels of CD (see Ize and Levy-Yeyati, 2003; Levy-Yeyati, 2006; Neanidis and Savva, 2009). Therefore, a plausible explanation of why high volatility leads to higher levels of CD is because it raises the expected cost of holding domestic-currency denominated liabilities.

Table 2 also displays a likelihood ratio (LR) test for the equality of mean values in short-run CD between regimes and its associated p-value. It reveals that the hypothesis of equal means is rejected at the 5% level of significance. For this reason, we proceed in our analysis with the model that allows for the expected component of short-term CD to vary by the regime of the common shock, as indicated in equation (6) above.

3.2 Test for shift contagion

In testing for shift contagion we examine the stability of the transmission of common shocks between low- and high-volatility regimes for our two country groups. For a preliminary visual inspection, Figure 2 presents the filtered probability of the common shock being in the high-volatility regime. There are pronounced and persistent periods of high volatility in the common shock, especially in the early part of our sample up to, and including, 1996. This could be due to the early abnormal experience of transition years where participants in the foreign currency markets were more favourable to foreign currency holdings because of the uncertainty that surrounded the success of market-oriented policies. There are also spikes in volatility in the second parts of 1997, 1998 and 2000 during the Asian, Russian and Argentinean crises, respectively. Periods of high volatility are also observed in 2002-2003, followed by a tranquil period of low volatility likely due to the full EU membership of eight countries in our sample (Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovak Republic and Slovenia) on May 2004 and of two further countries (Bulgaria and Romania) on January 2007.[22] Toward the end of our sample period some volatility has returned, probably a reaction to the global financial crisis.

For a formal test of shift contagion and estimates of the impact coefficients of the common shocks, we turn to Table 3. The impact coefficients of the common shock in the low-volatility regime, σ_{CH} and σ_{CL}, are small in size and hardly significant. Meanwhile,

[22] Neanidis (2010) indicates that the low CD volatility associated with EU membership is an outcome of the prospect of monetary integration with the euro area, which reduces fears of currency risk and raises the private sector's confidence in exchange rate stability.

their high-volatility counterparts, σ^*_{CH} and σ^*_{CL}, are much larger and statistically significant. Thus, for both country groups, the response to the high-volatility common shock is larger than that of the low-volatility shock. In other words, both groups display greater sensitivity to larger shocks. Moreover, the response of the LCD countries in the high-volatility regime is much more pronounced than the response of the HCD countries. This implies that LCD countries perceive domestic-currency denominated loans as having a higher risk component compared to HCD countries, leading them to switch into foreign-currency denominated credits by a greater margin. The implication of these results is that common shocks may have similar effects on the two country groups in 'normal' times, but disparate effects during turmoil periods. This is consistent with the notion of shift contagion.

We perform the statistical test for shift contagion and report its γ statistic along with its LR test statistic and the p-value. In the absence of shift contagion, the structural transmission of the common shock is stable across regimes in which case γ would be equal to one. Our results, however, reveal a value of γ different from unity, significant at the 1% level. This implies the presence of shift contagion between HCD and LCD countries. That is, the mechanism by which the common shock is transmitted between these country groups differs across volatility regimes. In other words, a large unexpected common shock to short-run CD affects countries differently depending on how heavily dollarized they are. But which country group exhibits greater reaction during the high-volatility common shocks? The value of γ being statistically smaller than one, dictates that there is a proportionately stronger reaction from the LCD countries. This higher degree of sensitivity to common shocks by the low-dollarized countries is in line with the findings of NS who show that spillovers in CD from within a region are much higher in low-dollarized environments. A plausible explanation has to do with the reaction of international banks operating in the region. In their effort to disperse the higher currency risk they are exposed to due to a common shock, they expand their portfolio of loans in foreign currency to countries with low levels of such assets. In this way, the banks' optimal reaction to rebalancing their assets internationally leads to a heterogeneous impact of a common shock in countries that differ in their holdings of foreign currency loans.

3.3 Tests for pure contagion

Tests for pure contagion investigate whether an idiosyncratic shock arising to one country group spills over to the other country group via channels that only exist during turbulent periods. Having already controlled for common factors, idiosyncratic shocks represent pure country group shocks. Since we cannot *a priori* justify the direction of contagion running in one way, we allow for the possibility of simultaneous bi-directional pure contagion. We first present the filtered probabilities of the idiosyncratic shock being in the high-volatility regime for each of the country groupings and, then, report the estimates for the impact coefficients of these shocks and the statistical tests for pure contagion.

Figure 3a illustrates the filtered probability of being in the turbulent state for the HCD countries, while Figure 3b does the same for the LCD economies. The graphs show that the idiosyncratic shocks for both country groups are located in the high-volatility state with high frequency until 1998. Post-1998, the short-term CD shock continues to be persistent for the HCD countries while it dissipates for the LCD countries. In total, HCD countries spend a larger proportion of time in the high-volatility regime compared to LCD countries. This marked difference in the behavior of the highly-dollarized countries suggests the higher idiosyncratic risk they exhibit compared to the low-dollarized countries. Both idiosyncratic shocks, however, appear to be widespread enough to cause concerns for spillovers across country groups.

Table 4 provides more information about the effects of the idiosyncratic shocks. Some notable features are as follows. First, the impact coefficients of the idiosyncratic shocks are highly significant at both the low- and the high-volatility regimes. Second, as for the common shocks, both country groups are more responsive to a high-volatility idiosyncratic shock with the response of the LCD countries being far greater than that of the HCD countries. Specifically, although in the HCD countries the estimated impact rises by about six-fold, in the LCD countries it increases by more than ten-fold. Finally, when compared to the impact coefficients of the common shocks, those of the idiosyncratic shocks are much larger in magnitude. This implies that when faced with similarly sized common and idiosyncratic shocks, both sets of countries are affected more by the latter. This supports the argument that a country's dollarization experience is mainly shaped and influenced by its own individual characteristics. Kishor and Neanidis (2015) confirm this by applying variance decomposition analysis, for every country in this study's sample, and find that the relative contribution of the idiosyncratic component to changes in CD is very high, on average 70%.

The coefficients δ_H and δ_L capture the presence, strength and direction of pure contagion between the two country groups. The strong statistical significance in each case suggests that high-volatility idiosyncratic shocks in short-run CD cannot be self-contained during episodes of heightened sensitivity but spread over to other countries. In other words, the transmission of the idiosyncratic shock is unstable between regimes so that during turbulent periods these shocks spill over from one country group to the other. Moreover, the significance of both δ_H and δ_L confirm the importance of modelling contagion effects as a bi-directional phenomenon whereby the transmission of the idiosyncratic shock runs from the high-dollarized to the low-dollarized countries and vice-versa.

Having established the presence of spillover effects in short-term CD, the negative signs of δ_H and δ_L reveal that a positive shock in short-run CD in either country group increases short-run CD in the other country group. Simply put, if either country group experiences a higher change in CD, this is transmitted to the other country group. This corresponds to pure bi-directional contagion in short-term CD across countries. But there is more to this finding. Table 4 also conveys information about the relative strength of pure contagion between the two country groups. A likelihood ratio test for the equivalence of the pure contagion coefficients is rejected implying that, with $|\delta_H| < |\delta_L|$, the highly-dollarized countries exert a stronger degree of pure contagion to the LCD countries. This result echoes that identified for shift contagion ($\gamma < 1$) and reveals that LCD countries are more sensitive to both common and idiosyncratic shocks, leading them to react to a greater degree in response to either shock compared to HCD countries. The difference in responses between the two country groups may be driven by the financial intermediaries' optimal realignment of their currency portfolios toward foreign currency loans internationally in an effort to minimize risks and losses associated with loan provision denominated in domestic currency.

The analysis so far established that countries where loans are dominated by the local currency are exposed to greater CD spillovers from the other countries in the region. However, conducting the analysis at the country group level may mask differences in the way countries within each group react to shocks in short-run CD. Following from this observation with so many different countries in each group, it is not clear what the idiosyncratic shock captures—it may be proxying a common shock at a lower aggregation level. Furthermore, the analysis thus far does not allow exploring the linkages between countries that give rise to contagion effects, especially since there exists only one coefficient estimate for contagion. For all these reasons, it is important to examine the robustness of our findings and repeat the analysis by pairing every country from the HCD group to every country from the LCD group. This exercise allows generating estimates of contagion for 132 pairs of countries which we can assess against the results offered by the country groupings. These estimates also allow testing explicitly for the channels of transmission of short-term CD across individual country pairs.

Figures 4a and 4b present the pairwise estimates of pure contagion for δ_H and δ_L respectively. Every point represents an estimated coefficient of pure contagion accompanied by the size of its respective standard error. Estimates are sorted by size from negative to positive and exclude in each case one outlier coefficient estimate, with values of 49.3 for δ_H and 20.8 for δ_L. The remaining 131 estimates have average values of -0.473 for δ_H and -1.035 for δ_L, indicated on the graphs with horizontal black lines. The plots also show at the horizontal axes the min and max values of the coefficients. The sign and magnitude of the average values corroborate the findings based on country groupings: on average there exists bi-directional pure contagion with its impact being greater in low-dollarized countries.[23] These results suggest that treating countries in groups is a good approximation for the cross-country pair variation in pure contagion. In simple terms, the analysis supports the argument that there are CD spillovers *across* countries that belong to groups that have widely different dollarization experiences. This complements NS who show the existence of regional CD spillovers *within* countries that

[23] A test that examines the statistical equivalence between the two mean values rejects the null hypothesis that these parameters are equal.

belong to groups that share similar dollarization experiences. We next use the coefficient estimates to tackle the identification of cross-country contagion effects.

3.4 The economic identification of pure contagion

Thus far, we have derived the estimates for shift and pure contagion from a purely statistical procedure without any economic input. It is useful, however, to understand why shocks transmit between countries that differ in their levels of credit dollarization. Doing so corresponds to the second step of our EDV approach that helps identify the contributing factors of pure contagion. Unlike the time-series orientation of the first step, the analysis in this step utilizes the cross-sectional variation between pairs of countries. This means that the dependent variable varies across pairs of countries but not across time, so that we take average values of all control variables over the entire time period. In this way, we estimate pure contagion against the time period mean value of each control variable.

Table 5 presents estimates of models explaining the variation in country-pair pure contagion as a function of geographical characteristics, a shared institutional environment, and macroeconomic links between countries. The regressions correspond to equation (9) of our econometric specification estimated with FGLS, where we correct standard errors following Hanushek (1974). We add control variables as we move to the right of the table, starting with geographical proximity. Columns (1)-(3) estimate the drivers of δ_H while columns (4)-(6) those of δ_L.

Pure contagion in both directions is explained by some common factors, notably adjacency and distance. Countries that share a common border and are closer to each other have higher spillovers in CD that run both ways, from HCD to LCD countries and from LCD to HCD countries. Bidirectional propagation of contagion is also explained by two macroeconomic indicators, the degree by which countries trade with each other and the economic size of the country where contagion originates from. The former effect is indicative of how interconnected countries are with respect to their trade transactions, while the latter is characteristic of a gravity effect. At the same time, countries differ with respect to the effect of banking penetration. It is only contagion running from HCD to LCD countries that increases with a greater exposure of HCD countries' banks to the international banking sector. When banks from LCD countries increase their share of foreign ownership, this does not contribute to pure contagion effects to HCD countries. The difference in the effect of foreign bank penetration on pure contagion contingent on the type of country (HCD vs. LCD) contagion stems from lends support to our findings in the first step of the analysis that LCD countries react more to pure contagion than HCD countries, i.e., that $|\delta_H| < |\delta_L|$. In other words, pure contagion from HCD to LCD countries is driven by the way international banks in the HCD countries reallocate their portfolios toward foreign currency assets in countries with low levels of such assets, not the other way around.

The final set of controls, those of EU membership, also has different effects on contagion depending where contagion starts from. Convergence to the institutional and operational environment of the EU encourages contagion from the HCD to the LCD countries ($\lambda_1 < 0$ and $\lambda_2 < 0$), although by a smaller margin when both countries within a pair are EU members ($\lambda_3 > 0$). This is in contrast to contagion effects spilling over from LCD to HCD countries which decrease in size when the HCD country is an EU member. These results extend the finding of Neanidis (2010) who shows that EU membership raises foreign currency lending within countries, to also encouraging foreign currency lending across countries. This cross-country spillover of credit dollarization appears to be in one direction, from HCD to LCD countries, giving rise to convergence in loan dollarization within the EU: countries with low CD levels experience a higher growth rate of foreign currency lending.

In an effort to quantify the importance of the factors contributing to pure contagion, we use the estimates from columns (3) and (6) to examine the magnitude of the effect that arises from a one standard deviation change in each control variable at a time. We exclude from this exercise the two dummy variables, contiguity and EU membership, since both are binary in nature rendering the exercise meaningless. For pure contagion from LCD to HCD countries, a one standard deviation decrease in distance from its mean translates into a 12.71 percentage point increase in pure contagion. For equivalent increases in bilateral trade and GDP, the respective figures are 29 and 24.5 percentage points. Calculating the magnitudes for pure contagion from HCD to LCD countries, a one standard deviation decrease in distance from its mean translates into a 38 percentage point increase in pure contagion. The respective figures for bilateral trade, foreign bank penetration and GDP are 5, 10, and 9 percentage points. This exercise, therefore, reveals that although all factors contribute to greater pure contagion, they do so to a different degree. Contagion from LCD to HCD countries reacts mostly to economic factors, whereas contagion from HCD to LCD countries responds mainly to geographic criteria.

Overall, results based on the cross-sectional analysis provide a measure of confidence that our regime-switching findings based on the bivariate analysis have sound economic foundations. In this light, it is clearer why during periods of turbulence countries experience contagion effects in short-run CD.

4. Conclusion

This paper has sought to cast further light on the contagion effects of credit dollarization across countries. The importance of this phenomenon is well-documented for the smooth functioning of the financial sector and for its macroeconomic implications. What is less well-understood is the extent to which the process of credit dollarization is solely determined by country-specific factors or whether it might be reinforced through linkages across countries. This question first posed in a companion paper, by NS, documented regional spillovers in both deposit and credit dollarization. In this paper, our primary objective is to take the analysis further in an effort to examine heterogeneous contagion effects in credit dollarization between countries that have different levels of dollarization. We do this in two steps. First, by investigating the presence of shift and pure contagion in credit dollarization. Second, by exploring the channels via which contagion in credit dollarization spreads. The long and diverse experience with credit dollarization of Central and Eastern European transition economies makes these countries a natural choice to apply our analysis.

NS provide evidence that credit dollarization spreads across national borders and that the transmission mechanism is trade and banking linkages amongst countries. This is the first study to show that the process of credit dollarization is not independent across countries but it depends on its evolution in the broader geographic region. They also show that the effects of spillovers are heterogeneous across countries, whereby they are more pronounced in nations with lower levels of credit dollarization. The analysis, however, focused on the stock measures of dollarization and restricted itself to countries that share similar dollarization experiences, i.e., within low-dollarized and within high-dollarized country groups. In this way, NS did not examine short-run variations in credit dollarization or whether credit dollarization spills over from one country group to the other. Hence, the question whether there is convergence in short-term credit dollarization between high-dollarized and low-dollarized countries remains unanswered.

Building on the findings of our companion study, the empirical analysis in this paper provides strong evidence that short-run credit dollarization spills over across countries with very different dollarization profiles. Specifically, we find that flow credit dollarization exhibits both shift and pure contagion between high-dollarized and low-dollarized countries. A further finding is that the group with low levels of credit dollarization reacts proportionately more to both types of contagion. This offers support to the notion of convergence in credit dollarization between country clubs. Beyond this finding, a key aspect of our analysis is that we employ a technique that permits us to estimate the contributing factors to contagion. In addition to the banking and trade linkages identified by NS, we also find geographic closeness, EU membership and a country's economic size to matter for the spread of flow credit dollarization. Amongst these, geography is more important for spillovers to countries with low levels of credit dollarization, while bilateral trade and economic size matter more for spillovers to high-dollarized countries.

Finally, our results may have important policy implications. They suggest that countries characterized by geographic proximity are well served to work together in an effort to achieve their individual desired levels of credit dollarization. Dismissing the effects of spillovers may lead policymakers to find themselves in a position of having to face higher credit dollarization than anticipated. For this reason, a coordinated strategy amongst countries that takes into account contagion effects is likely to be more successful in allowing them to reach their targeted levels of credit dollarization compared to a strategy that views a country operating in isolation.

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Table 1: Summary statistics

	All countries	High CD countries	Low CD countries
Credit Dollarization			
Mean	0.4149	0.6126	0.2234
Standard deviation	0.2291	0.1336	0.1071
Maximum	0.9348	0.9348	0.4141
Minimum	0.0175	0.4150	0.0175
Observations	3653	1810	1843
Δ(Credit Dollarization)			
Mean	0.0010	0.0017	0.0002
Standard deviation	0.0138	0.0170	0.0097
Maximum	0.1733	0.1733	0.1484
Minimum	-0.2267	-0.2267	-0.0648
Skewness	1.1748	0.6514	2.8471
Kurtosis	50.22	39.57	43.54
Jarque-Bera	63.8479	63.8821	67.4533
Observations	3630	1803	1827

Notes: Credit dollarization (CD) is defined as foreign currency denominated credit to total credits of residents issued by resident banks. Δ(Credit Dollarization) is defined as the first difference in credit dollarization, by country. High (Low) CD countries are those for which their mean value of credit dollarization exceeds (falls below) the mean value of credit dollarization of the entire sample--0.4149. Skewness is defined as m_3 / s^3, where m_3 is the centred third moment of the data and s is the sample standard deviation. Kurtosis is defined as $(m_4 / s^4) - 3$, where m_4 is the centred fourth moment of the data.

Figure 1a: Countries with high credit dollarization (HCD)

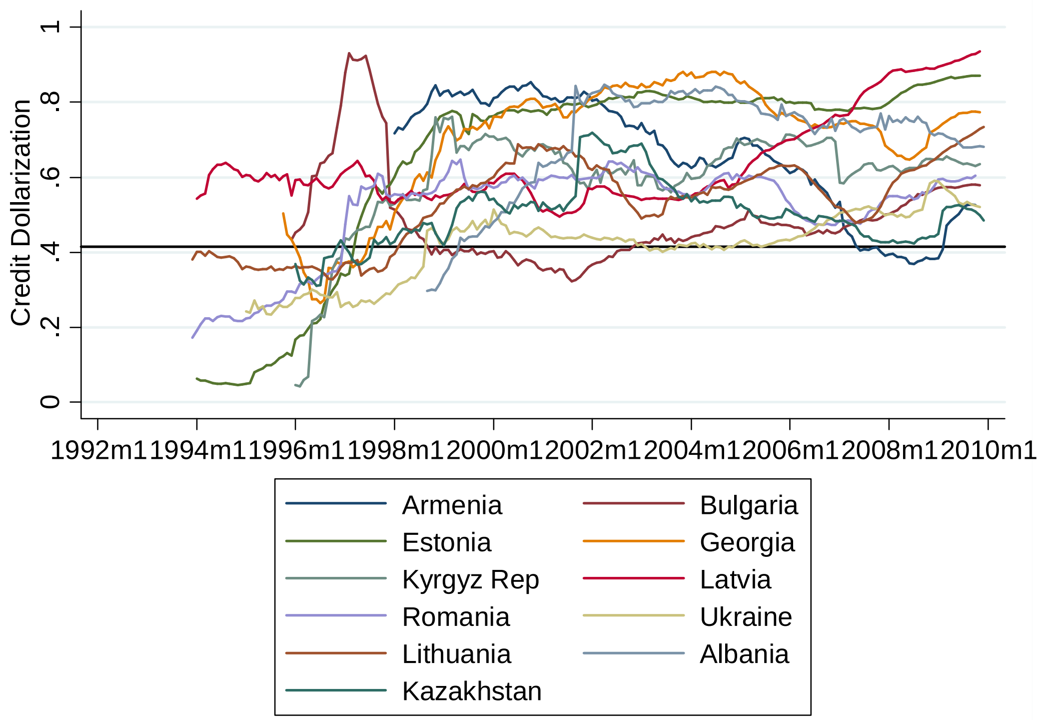

Figure 1b: Countries with low credit dollarization (LCD)

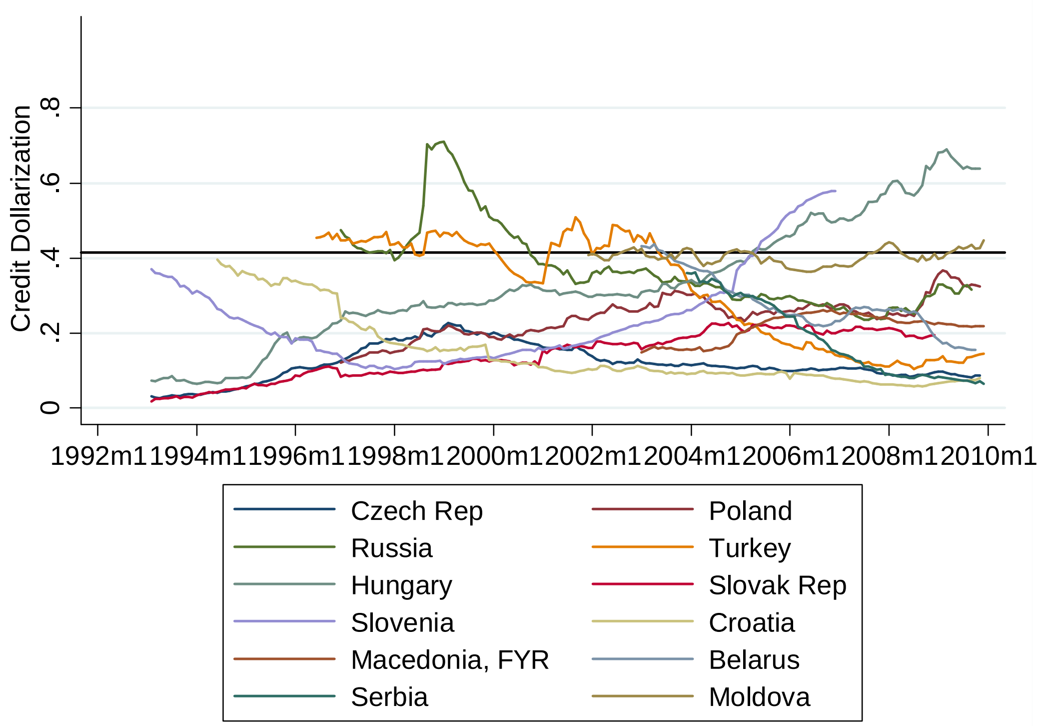

Figure 1c: First difference in credit dollarization (all countries)

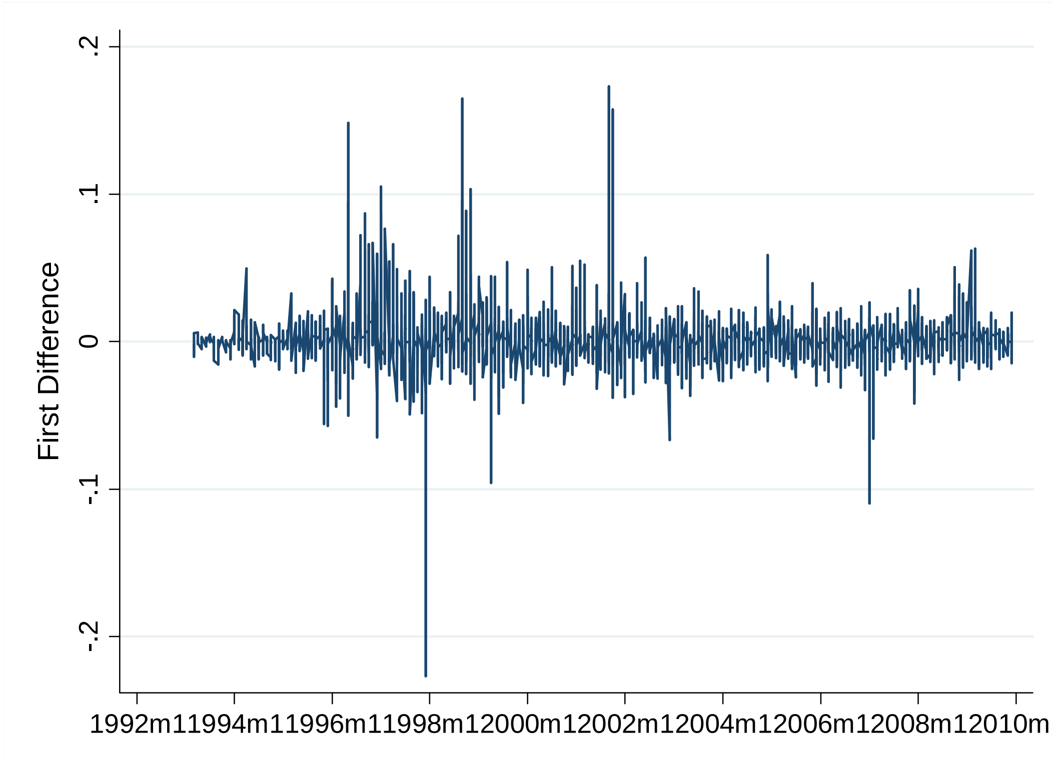

Table 2: Estimates of mean values in short-run credit dollarization across regimes

Threshold in credit dollarization	Low-volatility regime		High-volatility regime			
	μ_H	μ_L	μ_H^*	μ_L^*	**LR**	**p-value**
Mean (0.415)	-0.024 (0.051)	-0.087*** (0.028)	0.264*** (0.092)	-0.101 (0.071)	8.861**	(0.012)

Notes: μ_H and μ_L report the expected mean values of the first-difference in credit dollarization in the low-volatility regime for the HCD and LCD countries respectively, while μ_H^* and μ_L^* refer to the corresponding estimates for the high-volatility regime. Standard errors are reported in parentheses. The likelihood ratio (LR) statistic is for the null of equality of mean first-difference in credit dollarization across regimes. The test statistic has a Chi-square (2) distribution under the null hypothesis. ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively.

Figure 2: Filter probability of high-volatility common shock between HCD and LCD countries

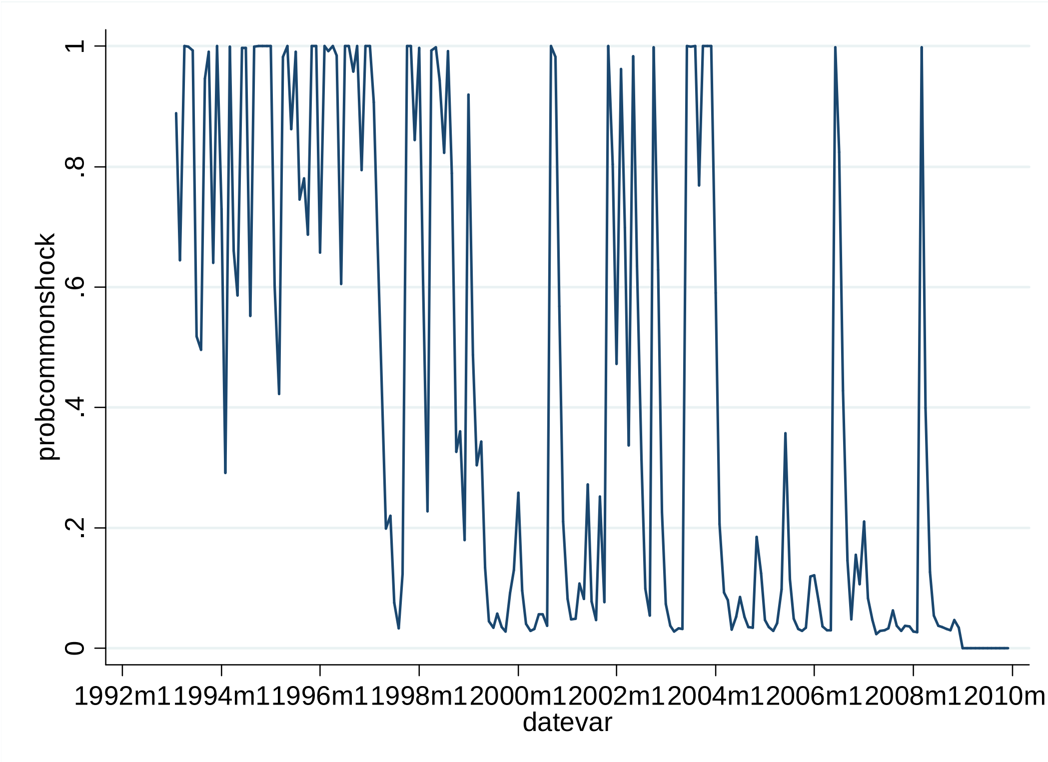

Table 3: Estimates of impact coefficients for common shocks

Threshold in credit dollarization	Low-volatility regime σ_{CH}	Low-volatility regime σ_{CL}	High-volatility regime σ^*_{CH}	High-volatility regime σ^*_{CL}	γ	LR	p-value
Mean (0.415)	0.111* (0.065)	0.043 (0.094)	0.219* (0.124)	0.452*** (0.118)	0.186	9.823***	(0.002)

Notes: σ_{CH} and σ_{CL} report the impact coefficients of common shocks for the first-difference in credit dollarization for the HCD and LCD countries respectively during the low-volatility regime, while σ^*_{CH} and σ^*_{CL} refer to the corresponding estimates during the high-volatility regime. Standard errors are reported in parentheses. γ is the ratio of responses to a common shock in the high-to-low volatility regime, as discussed in the text. The likelihood ratio (LR) statistic is for the null of no shift contagion against the alternative of shift contagion between HCD and LCD countries. The test statistic has a Chi-square (1) distribution under the null hypothesis. ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively.

Figure 3a: Filter probability of the HCD countries idiosyncratic shock being in the high-volatility regime

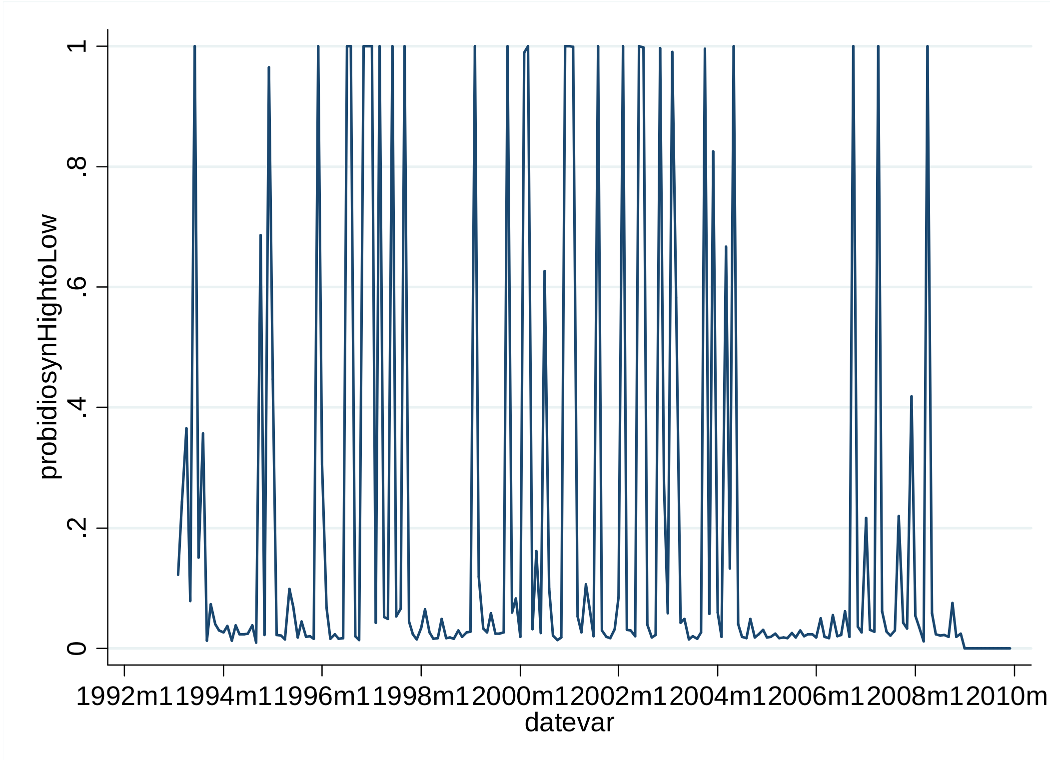

Figure 3b: Filter probability of the LCD countries idiosyncratic shock being in the high-volatility regime

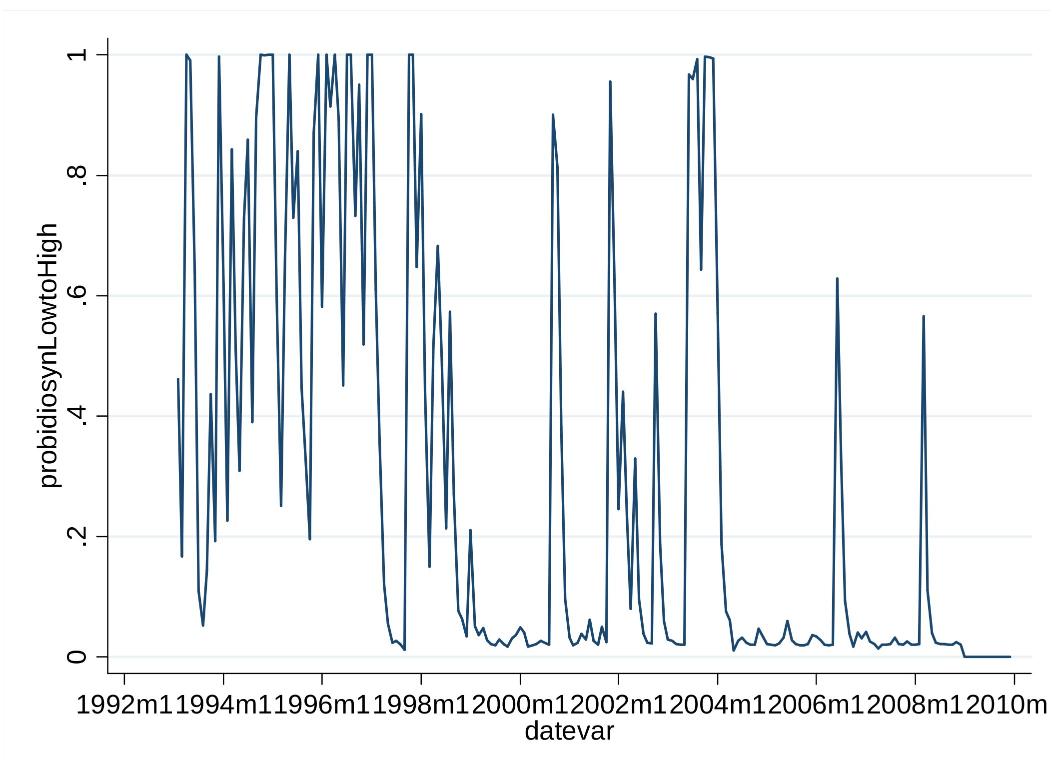

Table 4: Estimates of impact coefficients for idiosyncratic shocks

Threshold in credit dollarization	**Low-volatility regime**		**High-volatility regime**					
	σ_H	σ_L	σ_H^*	σ_L^*	δ_H	δ_L	**LR**	**p-value**
Mean (0.415)	0.287*** (0.045)	0.210*** (0.029)	1.559*** (0.378)	2.945*** (0.561)	-1.987*** (0.534)	-4.525*** (0.887)	10.421***	(0.001)

Notes: σ_H and σ_L report the impact coefficients of idiosyncratic shocks for the first-difference in loan dollarization for the HCD and LCD countries respectively during the low-volatility regime, while σ_H^* and σ_L^* refer to the corresponding estimates during the high-volatility regime. Standard errors are reported in parentheses. δ_H and δ_L capture spillovers during a high-volatility regime to the HCD and LCD countries, respectively. The likelihood ratio (LR) statistic is for the null of equal bi-directional pure contagion ($\delta_H = \delta_L$) against the alternative of different bi-directional pure contagion between HCD and LCD countries. The test statistic has a Chi-square (1) distribution under the null hypothesis. ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively.

Figure 4a: Estimates of pure contagion from LCD to HCD countries (δ_H), by country pair

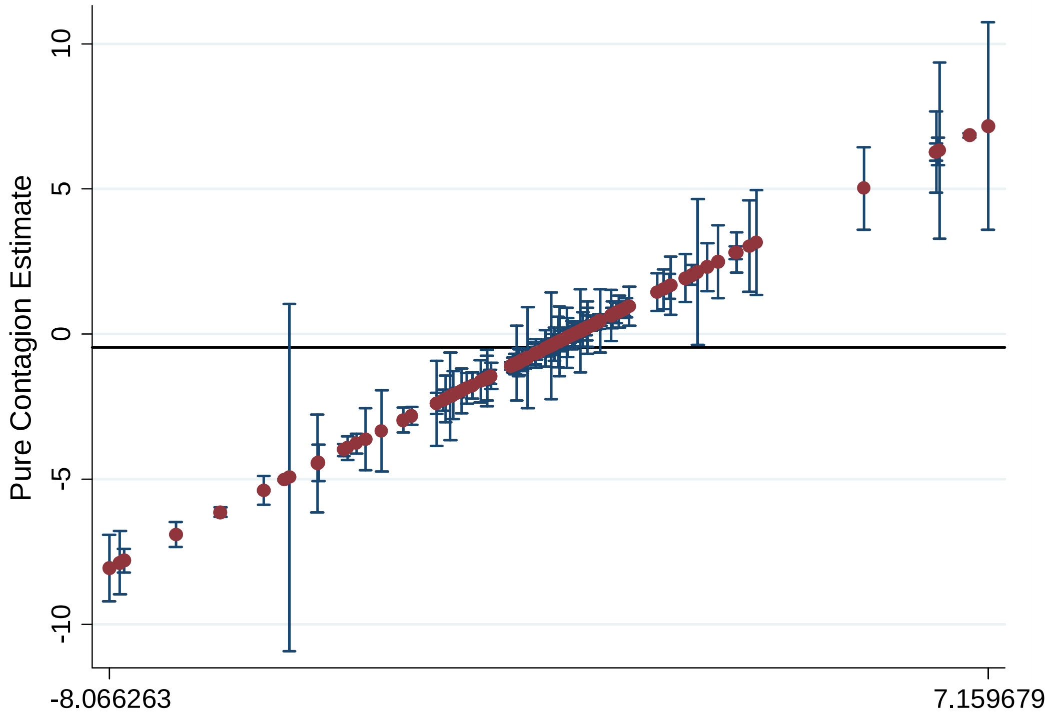

Figure 4b: Estimates of pure contagion from HCD to LCD countries (δ_L), by country pair

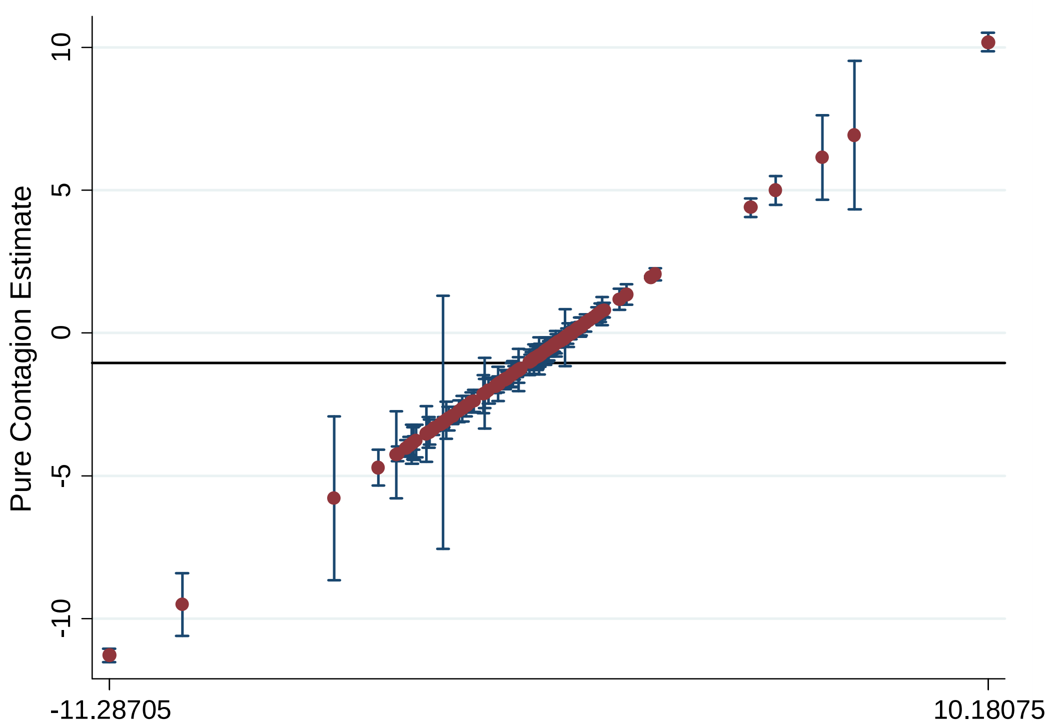

Table 5: Country-pair variation in pure contagion: geography, institutions and macroeconomic factors

Dependent variable →	**Pure Contagion from LCD to HCD countries (*δ_H*)**			**Pure Contagion from HCD to LCD countries (*δ_L*)**		
Explanatory variable ↓	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Contiguity	-0.193 (0.132)	-0.305*** (0.086)	-0.859*** (0.189)	-0.534*** (0.077)	-0.620*** (0.038)	-0.647*** (0.089)
Distance	0.0002*** (0.00001)	0.0001*** (0.00002)	0.0001*** (0.00003)	0.0003*** (0.00002)	0.0003*** (0.00002)	0.0003*** (0.00002)
EU member LCD		-0.007 (0.074)	0.057 (0.112)		-0.674*** (0.059)	-0.641*** (0.072)
EU member HCD		0.569*** (0.098)	0.441*** (0.155)		-0.498*** (0.106)	-0.285** (0.131)
EU member LCD & HCD		0.008 (0.099)	0.117 (0.158)		0.339*** (0.118)	0.299** (0.139)
Bilateral trade share			-8.54*** (2.28)			-0.392* (0.204)
Foreign bank penetration share			1.81 (1.29)			-4.70** (2.31)
Log GDP			-0.187*** (0.029)			-0.076** (0.030)
Obs	131	131	131	131	131	131

Notes: Explaining pure contagion estimates for all pairs of HCD and LCD countries with geography, institutional and macroeconomic variables. Dependent variable is the coefficient estimates of pure contagion for the country pairs estimated by our regime switching model and illustrated in Figures 4a and 4b. Standard errors are reported in parentheses based on generalized least squares estimations that account for cross-sectional heteroscedasticity and correct for bias in the standard error estimates following Hanushek (1974). Constant term not reported. Detailed explanation of explanatory variables and their sources appear in Table A1 in the Appendix. The sample period is February 1993-December 2009. ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively.

APPENDIX

Table A1: Credit dollarization data

Country	Credit Dollarization	Observations
Albania	1998:9-2010:1	137
Armenia	1998:1-2009:10	142
Belarus	2003:1-2009:10	82
Bulgaria	1995:12-2009:11	167
Croatia	1994:6-2009:11	186
Czech Rep.	1993:1-2009:11	203
Estonia	1993:1-2009:11	203
Georgia	1995:10-2009:11	170
Hungary	1993:1-2009:11	203
Kazakhstan	1996:1-2010:1	169
Kyrgyz Rep.	1996:1-2009:11	167
Latvia	1994:1-2009:11	191
Lithuania	1993:12-2009:12	193
Macedonia FYR	2003:1-2009:12	84
Moldova	2001:12-2010:1	98
Poland	1996:12-2009:11	156
Romania	1993:12-2009:11	192
Russia	1996:12-2009:9	154
Serbia	2003:12-2009:12	73
Slovak Rep.	1993:1-2008:12	192
Slovenia	1991:12-2006:12	181
Turkey	1996:6-2009:11	163
Ukraine	1995:1-2009:11	179

Table A2: Variable definitions and sources

Variable	Definition [source]
Credit dollarization	Foreign currency denominated credit to total credits of residents issued by resident banks [IFS and NCB]
Contiguity	Dummy variable that takes the value of 1 whenever a pair of countries share a common border and 0 otherwise. [Centre d'Études Prospectives et d'Informations Internationales (CEPII)]
Distance	Bilateral distance between two countries measured by the great-circle distance between the two economies' economic centers. [Centre d'Études Prospectives et d'Informations Internationales (CEPII)]
EU member LCD	Dummy variable for the LCD countries that takes the value of 1 during the dates of (i) the beginning of the EU accession negotiations, (ii) the end of the negotiation process, and (iii) after full membership to the EU, and 0 otherwise. [Author's calculation]
EU member HCD	As above but for the HCD countries [Author's calculation]
EU member LCD & HCD	A combination of the dummies "EU member LCD" and "EU member HCD".
Bilateral trade share	For total bilateral trade $\sum_{j=1}^{n} trade_{ij}$ of country *i* with the rest of the countries in the region and for bilateral trade $trade_{ij}$ with country *j*, averaging across all time periods *t*, the share of bilateral trade of country *i* with country *j* is $$Biltrade_{ij} = \frac{trade_{ij}}{\sum_{j=1}^{n} trade_{ij}}$$ [annual data on the value of merchandise exports and imports between each country and all its trading partners is from the IMF Direction of Trade Statistics (DOTS) database]
Foreign bank penetration share	For aggregate regional foreign bank penetration $\sum_{i=1}^{n} penetration_{it}$ of all countries at period *t* and penetration $penetration_{it}$ of country *i* at time *t*, averaging across all time periods *t*, the share of foreign bank penetration of country *i* is $$ForBankPenet_{it} = \frac{penetration_{it}}{\sum_{i=1}^{n} penetration_{it}}$$ [annual data on foreign ownership, defined as banks with assets of foreign ownership > 50%, are from the European Bank for Reconstruction and Development (EBRD) Banking Survey]
Log GDP	Logarithm of PPP GDP, defined as gross domestic product converted to international dollars using purchasing power parity rates. Data are in current international dollars. [WDI]