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Rural transformation, inequality, and the origins of microfinance☆

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ABSTRACT

What determines the development of rural financial markets? Starting from a simple theoretical framework, we derive the factors shaping the market entry of rural microfinance institutions across time and space. We provide empirical evidence for these determinants using the expansion of credit cooperatives in the 236 eastern counties of Prussia between 1852 and 1913. This setting is attractive as it provides a free market benchmark scenario without public ownership, subsidization, or direct regulatory intervention. Furthermore, we exploit features of our historical set-up to identify causal effects. The results show that declining agricultural staple prices, as a feature of structural transformation, leads to the emergence of credit cooperatives. Similarly, declining bank lending rates contribute to their rise. Low asset sizes and land inequality inhibit the regional spread of cooperatives, while ethnic heterogeneity has ambiguous effects. We also offer empirical evidence suggesting that credit cooperatives accelerated rural transformation by diversifying farm outputs.

1. Introduction

An influential literature stresses the role of finance for growth (King and Levine, 1993; Guiso et al., 2004; Lehmannn-Hasemeyer and Wahl, 2017; Heblich and Trew, 2018). However, we know little about the factors that originally gave rise to differences in financial development, and we know even less about the determinants of financial development over time.[1] Weak financial development is a particular concern in developing countries, especially in rural regions, where large segments of the population are unbanked, and access to financial services could potentially bring the greatest benefit (Burgess and Pande, 2005).

In this paper, we explain the rise and spread of micro finance institutions (MFIs), which are often seen as key institutions in providing financial services to rural areas (Armendariz and Morduch, 2010). Modern data show strikingly large differences in the spread of microfinance even within the same country.[2] Only a relatively small literature has attempted to explain this differential spread of small scale credit institutions across time and space (Rajan and Ramcharan, 2011; Ahlin et al., 2011; Vanroose, 2016; Colvin et al., 2018; Jaremski and Fishback, 2018). Although these contributions have added insight on the importance of individual variables, the choice of explanatory variables is often not informed by theory, and it is unclear how important individual variables are relative to each other.[3] Another fundamental challenge is simultaneity bias: MFI performance or market entry are observed at the same time as many of the explanatory variables, impeding causal inference. Moreover, as the universe of MFIs is rarely observable, and sampling is usually performed according to the quality of MFI

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[1] When attention to the origins of financial development is paid, the aforementioned literature often focuses on one time-invariant factor that can be used to instrument financial development.

[2] In India, for example, calculations on MIX Market data reveal the number of rural inhabitants per microfinance access point to differ substantially between districts. By this definition, outreach differs on average by a factor of 25 between the district with maximum and minimum coverage within the same state.

[3] An exception to the latter point is Ahlin et al. (2011), who do comprehensively consider the dynamic macroeconomic context in which MFIs expand.

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records, selection bias may loom large. Finally, many studies performing tests on modern data pool MFIs operating under different ownership structures, regulatory frameworks, and subsidy schemes (Morduch, 1999). Yet how would micro finance develop without public ownership, subsidization, and under a minimal regulatory environment?

We present what we believe to be as close as possible to such a free market benchmark scenario: the rapid rise and spread of credit cooperatives in 19th century rural Germany. These cooperatives provided small scale savings and loan services to previously unbanked populations. Crucially, they were owned by their members, did not receive subsidies, and were subject to a minimal regulatory framework, the changes in which we can control for (Guinnane, 2003). Moreover, because we observe the *de novo* creation of an entire system of microcredit, we observe most of our independent variables before the appearance of MFIs. This mitigates simultaneity bias. We further strengthen the causality of our arguments by using a range of instrumental variables.

Geographically, we focus on the eastern periphery of Germany before World War I, a setting that has a number of desirable features for our analysis. Its population was comparatively poor, rural, had restricted access to the emerging banking system of the time, and was marked by ethnic and economic inequality. Credit cooperatives thrived in this setting, being among the very first institutions world-wide to develop the principle of small-scale joint liability lending. As such, they influenced the development of micro finance in contemporary developing countries (Armendariz and Morduch, 2010). The microeconomic mechanisms that enabled credit cooperatives to successfully provide small-scale credit has also been the focus of extensive research (Banerjee et al., 1994; Ghatak and Guinnane, 1999; Guinnane, 2001). We are able to build on these contributions to formulate a simple model of the emergence of credit cooperatives.

The basic argument in this theoretical framework is as follows. Faced with a decline in the relative prices of agricultural staples, grain farmers have an incentive to switch to goods whose relative prices are rising (Fig. 1a and b). Switching production requires capital, but farmers are credit-constrained in the sense that they cannot access bank loans unless their private assets exceed a certain threshold. The possibility of pooling assets then provides the incentive to form credit cooperatives. Over time, therefore, the development of rural credit institutions will be determined by the relative output prices of agricultural goods (Fig. 1c). However, the emergence of cooperatives is constrained by their cost of capital, which decreases in group size and members' assets, but increases in members' ethno-linguistic heterogeneity, the bank lending rate, and economic inequality.[4] Variation in these factors essentially determines the regional spread of microfinance institutions.

We test these predictions using a newly collected data set containing the universe of 4941 credit cooperatives in the six eastern provinces of the Kingdom of Prussia, which was the largest state of the German Empire at that time. We use this data to construct a panel of 236 eastern Prussian counties, where the number of credit cooperatives entering the market per year at the county level is our main dependent variable.[5] Our analysis spans the entire period from the founding of the first credit cooperative in 1852 until 1913, the last year before the War. The results strongly support the predictions of the model: credit cooperatives are more likely to be founded in regions marked by large farms sizes and commercial banking infrastructure, and less likely to emerge where the land distribution is unequal. Over time, interest rates and relative grain prices are strong predictors of the development of the small scale credit sector. These results are robust to controlling for a wide range of variables, checking for simultaneity bias, and utilizing instrumental variable procedures. Finally, we show how our parsimonious set of predictors successfully explains a large share of the variation in cooperative credit institutions in both the cross section and time dimension.

We link the decline in relative grain prices we observe in the data to an influential literature stressing the role of income effects in the process of structural transformation (Kongsamut et al., 2001; Timmer, 2007; Herrendorf et al., 2014). This literature often shows how non-homothetic consumer preferences lead to a decline in the relative demand for agricultural staples as consumer incomes rise (Engel's Law). Moreover, in an open economy, the price of staples, which are to a higher degree tradeable, is also influenced by (potentially lower) world market prices. Declining relative demand and import competition thus lead to a decline in the relative price of basic agricultural goods against capital-intensive goods. This provides an incentive to exit agriculture, thus shifting labor into capital-intensive sectors (Uy et al., 2013; Teignier, 2018). We stress that a similar process was at work *within* 19th century agriculture. Declining prices for staple grains led farmers to exit that sector in favor of more capital-intensive meat and dairy production, a process we term "rural transformation". Apart from showing that relative grain prices crucially determined the creation of credit cooperatives, we provide suggestive quantitative evidence that these cooperatives indeed aided rural transformation. They did so by helping to set up production cooperatives that produced more capital-intensive agricultural products such as milk and meat (Fig. 1d). This conclusion is further supported by our examination of the Prussian cattle statistics. Such a mechanism is in line with the observations of qualitative economic historians (Kindleberger, 1951), but has to the best of our knowledge not been empirically investigated. As such, we provide a novel perspective that focuses on the role of MFIs in aiding transformation in developing rural regions.

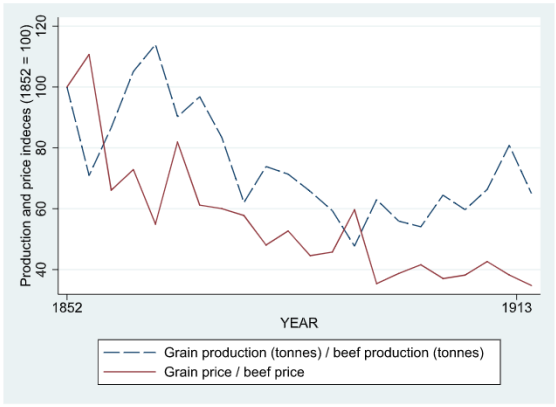
Our paper is also related to a long-standing debate in the micro finance literature regarding the existence of a tradeoff between outreach and financial sustainability of MFIs (Hermes et al., 2018; Cull et al., 2007). This literature focuses on the role of subsidies or regulation in affecting outreach and survival probability. While our measure of market entry of credit cooperatives is informative with regard to notions of outreach, we do not have anything to say about the survival of our cooperatives. This is both a feature of our data, as we do not observe failing institutions, and a feature of our historical setting: credit cooperatives in 19th century Germany had negligible failure rates (Banerjee et al., 1994). While this is an important caveat, it also provides us with a attractive empirical setting, as attrition is ruled out.

A further theme that is connected to our work is the market structure of the small-scale finance industry. In particular, the literature has analysed the interactions of micro-lenders and banks on market outcomes (Cull et al., 2014; Périlleux et al., 2016). We offer evidence to show that the relationship between both types of institutions is complementary, at least at the point of MFI entry. Intimately related to this is a strand of the literature that empirically analyzes the determinants of institutional features of MFIs, such as the liability structure (Banerjee et al., 1994; de Quidt et al., 2018). We primarily focus on the determinants of MFI entry over space and time, rather than focusing on its institutional structure. We do, however, show that changes in nationwide liability laws had a large effect on market entry decisions.

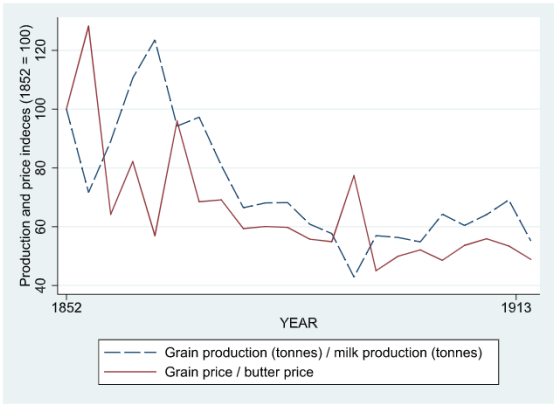
A final strand of literature we contribute to analyzes the effects of ethnic and economic heterogeneity on the participation in clubs. In a pioneering paper, Alesina and La Ferrara (2000) show that high levels of ethnic and income inequality are empirically associated with lower participation in social clubs in US cities, with ethnic heterogeneity being the most important driver. Similarly, La Ferrara (2002) finds that an unequal distribution of assets decreases club participation in Tanzanian villages, while Barr et al. (2015) find economic inequalities

[4] Once the number of cooperatives in a region is allowed to exceed 1, the prediction for ethno-linguistic heterogeneity is theoretically ambiguous, as heterogeneity can make it advantageous to split cooperatives according to ethnicity.

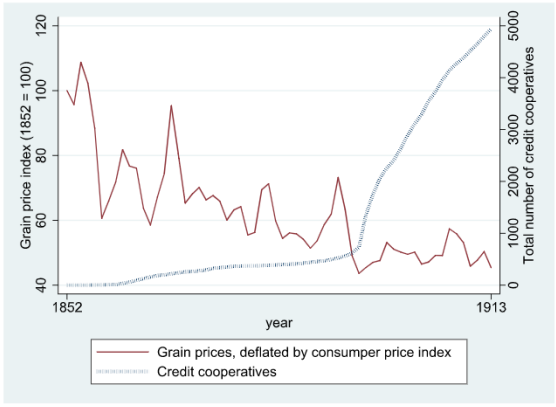
[5] We also have access to detailed balance sheet data for a small group of cooperatives.



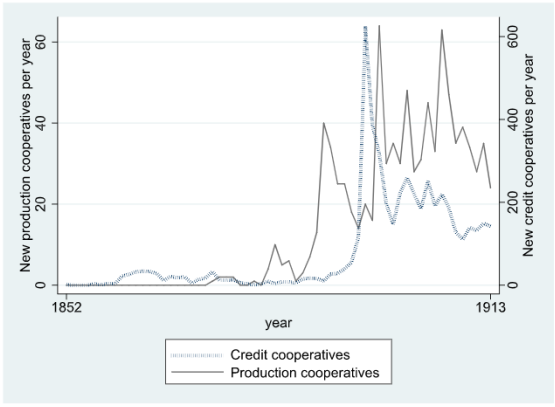
(a) Relative prices and production of grain and meat



(b) Relative prices and production of grain and milk



(c) Relative grain prices and credit cooperatives



(d) Credit and production cooperatives, changes

Fig. 1. Rural transformation in Prussia, 1852–1913. Six eastern provinces, grain refers to rye and wheat, weighted by their initial land shares. Source: production data from von Finckenstein (1960), other sources see text. Calculations: authors.

and ethnicity did not significantly affect the formation of community-based organizations in Zimbabwean villages. Similar arguments have been made less formally for the formation of production cooperatives in industrializing Europe. For example, O'Rourke (2007) argues that ethno-confessional fragmentation and an unequal distribution of land was decisive in holding back dairy cooperatives in Ireland. Fernandez (2014) finds some evidence that land inequality decreased the output produced by production cooperatives in Europe before the Second World War. Similarly to Alesina and La Ferrara (2000) and La Ferrara (2002) we model the formation of clubs, but we do this specifically for credit cooperatives, rather than social clubs. Furthermore, we stress that the formation of these clubs cannot be understood by solely looking at time-invariant heterogeneity, but that *changes* in economic conditions are key drivers of club formation.

The remainder of the paper is structured as follows. Section 2 introduces our conceptual framework, while section 3 outlines how we take the predictions of that framework to the data. Section 4 analyzes the results, and provides an assessment as to the relative importance of individual determinants of the growth of microcredit institutions. The final section concludes.

2. Conceptual framework

We model one region populated by N small agricultural producers. Producers are identical, with the exception of speaking different languages A, B so that $N_A + N_B = N$.[6] Producers choose between producing two goods: a land-intensive agricultural staple 'grain' and a capital-intensive output which we refer to as 'dairy'. Initially, farmers produce the agricultural staple only. For simplicity, we assume that each good uses only one specific factor of production:

$$\begin{aligned} Y_g &= Z^{\alpha} \\ Y_d &= K^{\beta} \end{aligned} \tag{1}$$

where Z are land inputs and K capital inputs.[7] Being small, each producer maximises profits given both output prices $\{P_g, P_d\}$ and input prices for land and capital $\{x, r\}$. Total profits are:

$$\begin{aligned} \Pi &= \pi_g + \pi_d \\ \pi_g &= P_g Z^{\alpha} - xZ \\ \pi_d &= P_d K^{\beta} - rK \end{aligned} \tag{2}$$

2.1. Cooperatives and the cost of capital

We are primarily interested in the procurement of capital. Farmers can receive a loan for the capital K that maximises profits π_d from a

[6] We further relax the assumption of identical producers below by allowing for asset inequalities.

[7] The assumption of specific factors eases the exposition of the model, but the results would carry through if we allowed capital to be mobile between sectors, as long as the production of the 'dairy' good is more capital intensive. This would echo the Hansen-Prescott approach to Unified Growth Theory, which seeks to model the structural transformation from an agricultural Malthusian to an industrial Solovian economy (Hansen and Prescott, 2002).

conventional bank at the rate of r_b against their private collateral q. However, we assume that banks are only willing to lend at the rate r_b if a borrower's collateral exceeds a certain threshold: $q \geq q_T$. This may reflect the fact that banks possess limited information on borrower quality, or face fixed monitoring costs, which makes dealing with lowly collaterized individuals unprofitable (Ghosh et al., 2000). The implication of this market imperfection will be that some farmers are credit constrained by a lack of collateral (Karlan and Morduch, 2009).

Producers for whom $q < q_T$ can overcome their lack of private collateral by pooling assets with other producers (Banerjee et al., 1994). Allowing $N_c \leq N$ farmers to join such a cooperative, the cooperative's total assets are then $Q = qN_c$.[8] We assume that the cost of capital r_c to each producer inside the cooperative decreases in both q and N_c. There are three motivations for this scale effect. Firstly, the cooperative may act as both a credit and savings institution, and extend loans to members primarily on internal deposits. In this case, increasing cooperative size may reduce the likelihood of negative shocks to individual members' assets (Guinnane, 1997). Secondly, pooled assets can be used by the cooperative as collateral against which to obtain loans from banks or other financial institutions. Assuming that bank rates are inversely proportional to the amount of collateral provided, pooling will lower the rate obtained. Moreover, given joint liability systems under which members guarantee loans reciprocally, more guarantors implies a higher likelihood that loans will be paid back (which may also improve access to third party financing) (Ahlin, 2015). Therefore, q and N can be thought of as substitutes: cooperatives can be smaller in wealthier regions where deposits are high, but large group sizes are required if average deposits are small.

However, the cooperative is only able to fulfill these functions if it has access to either local information on its members (Guinnane, 2001) or to an enforcement mechanism, such as joint liability lending (Ghatak and Guinnane, 1999). Both may be hampered by existing ethno-linguistic, religious or other cultural differences in the population.[9] In our model, we focus on the extent to which gathering information on and monitoring members is more difficult in the presence of communication costs (Guinnane, 2001), which are likely to run along ethno-linguistic lines.[10] We therefore assume r_c to be increasing in a cooperative's ethno-linguistic heterogeneity H_c, defined as:

$$H_c = 1 - \left| \frac{N_{A,c}}{N_c} - \frac{N_{B,c}}{N_c} \right| \tag{3}$$

where $N_{A,c}$ is the number of club members that belong to ethno-linguistic group A. Note that $0 \leq H_c \leq 1$ by this definition. In sum, therefore, we have for the cost of capital in the credit cooperative:

$$r_c = f(q, N_c, H_c) \tag{4}$$

where $f_q(.) < 0, f_N(.) < 0, f_H(.) > 0$.

The cost of capital each producer faces is a decreasing function of private assets or collateral q, which by virtue of assumption is identical to the club's average collateral. The remainder of r_c is determined by the characteristics of the cooperative: the benefits of group size, and the costs of heterogeneity.

[8] In practice, roughly a third of cooperative loans were directly collateralized (by land), whereas much of the remainder was guaranteed by co-signers, who had to pledge land or other property, the value of which sometimes exceeded the amount of the loan. See historical appendix G, section G.4.

[9] Ghatak and Guinnane (1999, p. 196) refer to this as "(…) the general idea that people with connections of shared locality or other *bonds based on kinship* and occupation may be able to support credit contracts that would be impossible with conventional banking practices." (emphasis added).

[10] Alternatively, one could think of social capital Karlan (2007) being smaller between members of different ethno-linguistic groups.

2.2. Cooperatives and rural transformation

We now impose some structure on the determinants of r_c. One possible reformulation of equation (4) is as follows:

$$r_c = r_b + \frac{1}{\delta q N_c} + H_c \quad \forall \quad q < q_T \tag{5}$$

where $0 < \delta \leq 1$ is the degree to which a cooperative's total assets qN_c decrease capital costs. This explicit formulation is attractive for two reasons. Firstly, the parameter δ can reflect institutional factors such as the liability structure the co-op chooses, which may be dependent on the legal framework (Banerjee et al., 1994). Changes in the legal framework would therefore influence the cost of capital of all cooperatives through this parameter. Secondly, equation (5) embodies the idea that at the limit, if cooperative membership is homogenous ($H_c = 0$), the cost of capital approaches the bank rate as club size N_c becomes very large.[11]

What factors will determine the emergence of cooperatives? Farmers for whom $q < q_T$ will found a credit cooperative to enter capital intensive production and exit grain production if $\pi_d \geq \pi_g$. Substituting 5 into 2 yields:

$$P_d K^{\beta} - (r_b + \frac{1}{\delta q N_c} + H_c)K \geq P_g Z^{\alpha} - xZ \tag{6}$$

Under the (temporary) assumption that only one cooperative can be founded, the region's demographics will mirror cooperative membership in equation (6) ($N = N_c, H = H_c$).[12] Simple comparative statics then indicate that:

PROPOSITION 1.

1. A decline in relative grain prices (P_g/P_d) shifts resources to the capital-intensive sector and will therefore lead to an increase in demand for cooperative credit. This mechanism is reflected in the pattern in Fig. 1 in the introduction, and predicts an important role for cooperatives in driving rural transformation.
2. A decrease in bank interest rates r_b or an increase in the pool of members N_c increases the likelihood of cooperatives being founded.
3. Low levels of private assets q will prevent farmers from assembling enough assets in a cooperative. Cooperatives are therefore more likely to appear in wealthier regions.[13] However, this is only true as long as $q < q_T$, as otherwise producers would turn to bank rather than cooperative credit. In sum, we would expect a hump-shaped relationship between q and co-op emergence.

2.3. Heterogeneity and endogenous number of co-ops

According to equation (6), a very heterogeneous population can prohibit the shift to Y_d and deter cooperative formation. This predic-

[11] Using micro-data on 135 credit cooperatives for which balance sheet information is available from police records (Polizei Präsidium Posen, 1909), we have also tested whether the relationships implied in equation (5) hold empirically. The results show a negative, though not always statistically significant, relationship between the cost of capital at the co-op level and capital stock per member, as well as between the cost of capital and the total number of members. We do not observe heterogeneity at the co-op level, nor bank rates in this particular data set.

[12] H is the heterogeneity in the regional population, defined analogously to equation (3).

[13] Even today, collateral still features heavily in microfinance. For example, Banerjee et al. (2015) report that half of surveyed lenders requested some collateral from borrowers in modern settings, in some cases in contravention of lender's guidelines. Note that cooperatives in the 19th century also demanded a capital contribution from members. In our data, the minimum size of this contribution is 10 Mark, roughly equivalent to 2 weeks' wages for agricultural laborers in the poorer counties.

tion would be in line with the results of the empirical literature stressing the difficulties of cooperatives to form in heterogeneous societies (O'Rourke, 2007). However, as the theoretical literature on fragmentation and club participation (Alesina and La Ferrara, 2000) stresses, this prediction no longer holds once the number of clubs is allowed to exceed 1. In this case, heterogeneity presents an incentive to segregate by ethnicity, which can lead to *more* clubs being founded. In particular, it can be shown that forming two (ethnically segregated) rather than one (ethnically mixed) cooperative lowers the cost of capital under certain conditions:

PROPOSITION 2. Founding two rather than one cooperative will minimize r_c if:

$$qH_N N > \frac{1}{\delta}$$

Therefore, the number of cooperatives can increase in regional population heterogeneity.

PROOF. Appendix. The interpretation is straightforward. The higher heterogeneity in the population, the more costly one mixed club becomes, and the more attractive it is to found two separate cooperatives. Differently from before, the number of cooperatives will now *increase* in population heterogeneity. The relationship between heterogeneity and the number of cooperatives is therefore theoretically ambiguous.

The other variables in PROPOSITION 2 work in the same direction as in PROPOSITION 1: the larger the population, the more viable two segregated (and therefore smaller) clubs will be. Finally, higher levels of collateral in a region will decrease the need for producers to pool resources, so two smaller segregated clubs are more efficient.

2.4. Inequality

We now allow private assets to differ within the region's population. For simplicity, we assume that individual assets follow a uniform distribution $q_i \sim U(\underline{q}, \overline{q})$. We let the mean of the distribution be equal to the regional mean used in the previous sections ($q = \frac{(\underline{q}+\overline{q})}{2}$). Given constant population N, this also implies that total assets in the region are unchanged. However, inequality can now affect co-op formation through its influence on a cooperative's assets, which can be expressed as:

$$Q = \int_{\underline{q}}^{\overline{q}} \left(\frac{N}{\overline{q} - \underline{q}} \right) dq - \int_{q_T}^{\overline{q}} \left(\frac{N}{\overline{q} - \underline{q}} \right) dq \tag{7}$$

Conditional on mean assets being below the threshold for bank financing, higher asset inequality implies an increasing $\overline{q}$ in equation (7). This increases the size of the second term, and thus decreases cooperative assets. As we know that co-op formation increases with co-op assets from PROPOSITION 1, it follows that:

PROPOSITION 3. For constant mean assets, asset inequality decreases credit cooperative formation conditional on $q < q_T$. Intuitively, inequality pushes a larger fraction of the population over the collateral barrier q_T required to obtain formal bank financing, and therefore decreases the total assets available for pooling to the remaining co-op members.[14] The effect of inequality will be even larger, if in addition, a minimum level of assets are required for participation due to mandatory capital contributions. If the mean q is held constant, an increasing $\overline{q}$ implies a falling $\underline{q}$, and therefore increases the fraction of the population who will not qualify for participating in the club.[15]

In summary then, we expect that credit cooperatives form in response to falling relative grain prices, as well as low bank interest rates. Moreover, we expect to find more credit cooperatives in regions with lower land inequality and higher private assets, although this latter relationship is likely to be hump-shaped. The relationship with ethnolinguistic heterogeneity is theoretically ambiguous.

3. Data and empirical strategy

3.1. Sample: Prussia's six eastern provinces

Our sample comprises the 236 counties that made up the six eastern provinces of the Kingdom of Prussia in the half-century before World War I. Fig. 2a shows the location of these provinces within the Kingdom of Prussia, which from 1871 onward was part of the German Empire.[16] These six provinces are often grouped together on account of their geographic location east of the river Elbe, and their joint history as destinations of German settlement and military expansion since the High Middle Ages, thus forming a coherent sample for analysis.[17]

Geography and this settlement history resulted in the creation of large landed estates and a predominantly rural orientation of the economy that marked the six provinces well into the 20th century (Cinirella and Hornung, 2016). Even in 1883, 50% of arable land was devoted to growing staple grains. However, rural transformation was gathering pace from the mid-19th century onwards. Berlin, one of the fastest growing urban agglomerations of the time, was located within Brandenburg, the westernmost of the six provinces. Urban demand for foodstuffs, especially dairy, meats, vegetables and alcohols, was expanding (Burhop, 2011). By rural standards, these were capital-intensive sectors.[18] At the same time, the import competition from land-abundant grain producers in the New World was intensifying, as advances in shipping technology outpaced protectionist measures by German policy makers (Findlay and O'Rourke, 2009). The consequence of increasing relative demand for non-staple foods, and of import competition in grains, was the downwards trend in grain prices identified in Fig. 1.

The medieval proliferation of landed estates in the six provinces had resulted in high levels of land-, and hence wealth inequality. Inequality had persisted despite the official liberation of the peasants in 1807 and despite the fact that many had become independent freeholders by the 1850s (Eddie, 2008).

Demographically, Prussia's eastern provinces were highly diverse. Some had been annexed by Prussia from the Polish Commonwealth in the late 18th century and therefore contained large Polish-speaking populations. Ethno-linguistic cleavages between Poles and Germans also largely correlated with religious divisions, with Poles being almost

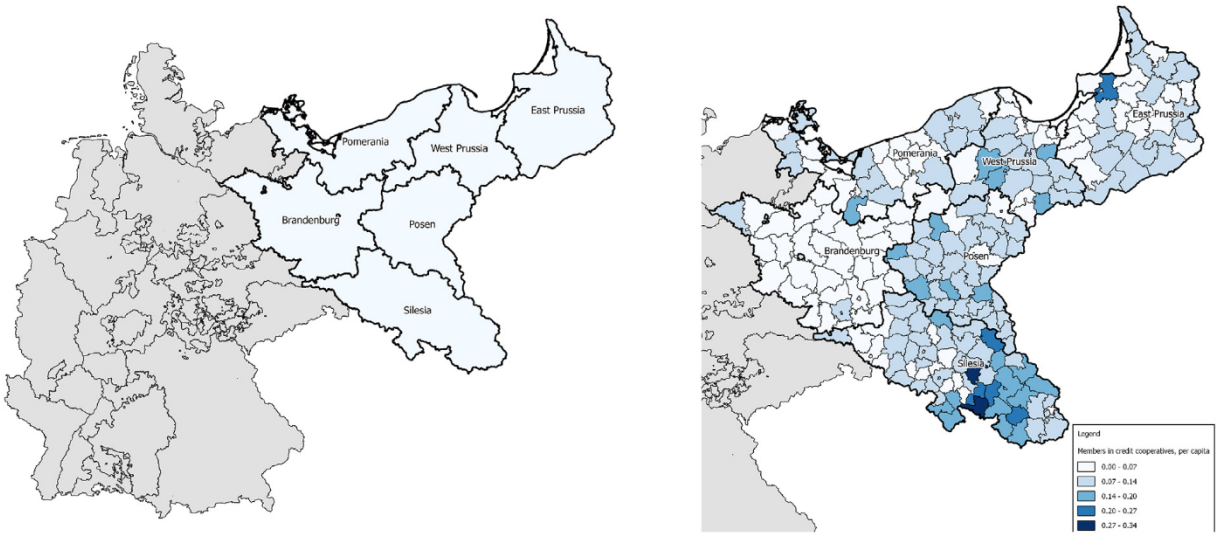
[14] This prediction is in line with La Ferrara (2002), who shows that rising asset inequality leads to less participation in clubs in developing countries, because the wealthy have less to gain from joining.

[15] For most of the analysis, we interpret higher inequality as implying a larger mass of assets in the upper tail of the distribution, although we do relax this interpretation to investigate the entire asset distribution in empirical section 2.

[16] Each province was subdivided into two or three districts (*Regierungsbezirke*), most of which encompassed between 12 and 20 counties. See table D1 for the administrative divisions of Prussia.

[17] In German historiography, these provinces are often jointly referred to as "East Elbia", a categorization that goes back to the writings of Max Weber (1892). See historical appendix, section G.1.

[18] O'Rourke (2007) mentions dairy creameries in Denmark requiring the milk of at least 400 cows to profitably operate technological innovations such as the cream separator. Procuring cattle was a significant expense for rural populations in Prussia too: In 1883, the price of a single head of beef cattle stood at roughly 187 Mark. Evaluated at the average rural wage rate, this translated into roughly 149 days of labor (See historical appendix, section G.2 for more information).



(a) The six eastern provinces of Prussia (b) Credit cooperative members per capita, 1913

Fig. 2. Provinces, counties and cooperatives in the sample.

exclusively Catholic, and Germans predominantly Protestant.[19] Jews constituted a third religious group in some regions, although their mean share had fallen to below 1% by the outbreak of World War I due to emigration.

Finally, the six provinces witnessed an unprecedented boom in the foundation of credit cooperatives. Credit cooperatives originated in the early 1850s following the activities of two pioneers, Friedrich Wilhelm Raiffeisen and Hermann Schulze-Delitzsch in western regions of Germany.[20] While Schulze-Delitzsch focused on urban craftsmen, and Raiffeisen on rural populations, both emphasized the principle of joint-liability lending to overcome the credit constraints of previously unbanked populations (Faust, 1965). Although membership was required to receive a loan from a credit cooperative, the capital contribution required for membership was relatively small and the co-ops savings facilities were often available for non-members too (Banerjee et al., 1994). Membership was open to all by residents of a locality and locals constituted the majority of directors and officers of the co-op. As a cooperative only operated in a closely circumscribed locality, and meetings were held in public, lenders and borrowers knew each other (Guinnane, 2001).[21]

As is visible from Fig. 2b, credit cooperatives were highly successful in attracting members, with close to 10% of the population being members at the end of our sample period in 1913. Most importantly for our purposes, there is extensive variation in the spread of cooperatives even between adjacent counties.

3.2. Dependent variable: credit cooperatives

The main data source on credit cooperatives is the 'Address Book of Purchasing and Economic Cooperatives in the German Empire' as produced by the central Prussian co-operative bank (Preußische Central-Genossenschafts-Kasse, 1915). This publication, which to the best of our knowledge has not been used before, provides the location and type of each cooperative in Imperial Germany, the year of its foundation, the number of members, the sum for which its members are liable, and the association to which it adheres. We use this data to construct a panel consisting of the number of credit cooperatives for each county from 1852 to 1913. This produces the data plotted in Fig. 1 and E.1a. As we observe only cooperatives active in 1915, we do not observe failing cooperatives. Fortunately, our theory concerns the foundation of cooperatives, rather than their demise and apart from a brief episode in the 1880s, failure rates of credit cooperatives were below 1 in 1000 (Banerjee et al., 1994; Guinnane, 2001). Nonetheless, this limitation of our data implies we are not in a position to investigate cyclical patterns of booms and busts in cooperative operation, and focus on long-term growth in the number of cooperatives instead.

We use the same source to construct a panel of non-financial cooperatives, which we use as a robustness check. The most important of these are production cooperatives, which were set up by farmers to jointly produce capital intensive goods such as dairy, meat, alcohol; that is those goods described by Y_d in our model.

3.3. Independent variables

3.3.1. Prices

The theory developed in section 2 predicts that declines in relative grain prices should increase cooperative credit. We use data on German rye and wheat prices from the official Prussian statistics (Preußisches Statistisches Landesamt, 1907; 1925). Rye and wheat are the staple grains most affected by import competition and changing urban demand, and are thus most suited to studying rural transformation.[22] We deflate German rye and wheat prices by the Germany-wide consumer price index calculated by Kuczynski (1961), so that we evaluate the evolution of grain prices relative to all other prices.[23]

A change in grain price defined in this way will be equal for all counties. However, exposure to this price shock will differ locally as the initial reliance on rye or wheat cultivation differs between coun-

[19] The correlation between the share of Catholics and Polish speakers at the county level is 0.72.

[20] The first credit cooperative in our sample was founded in 1852.

[21] More detailed information on the institutional structure of credit cooperatives is provided in the historical appendix, section G.4.

[22] Barley was produced only in relatively small quantities, while oats were partly used as an animal feed.

[23] The results are similar when using the price of a single good, such as butter, as a deflator (see column (6) in table F3). While this is closer to the price ratio P_g/P_d in the model's PROPOSITION 1, it does not capture the full range of goods that producers could switch to, which included not only dairy, but also meats, alcohols, and even proto-industrial textiles.

Table 1
Determinants of credit cooperative growth, county level, Prussia: **Benchmark**.

	(1) No fixed effects	(2) No urbanization	(3) Benchmark	(4) Controlling for trade & industrialization	(5) Controlling for ideas transmission	(6) Year fixed effects
	Dependent Variable: Number of new credit cooperatives per county					
Δ grain prices	−41.6120***	−38.2984***	−38.3988***	−38.3154***	−38.3955***	−38.3075**
	(3.2788)	(3.0438)	(3.0575)	(3.0454)	(3.0567)	(13.4610)
limited liability	1.6035***	1.5914***	1.5920***	1.5917***	1.5920***	
	(0.1547)	(0.1539)	(0.1540)	(0.1540)	(0.1540)	
central bank	1.3803***	1.3892***	1.3888***	1.3891***	1.3888***	
	(0.1419)	(0.1410)	(0.1411)	(0.1412)	(0.1411)	
incorporation	0.0821	0.0664	0.0667	0.0663	0.0667	
	(0.1313)	(0.1312)	(0.1313)	(0.1313)	(0.1313)	
population	0.0056***	0.0060***	0.0067***	0.0080***	0.0068***	0.0068***
	(0.0016)	(0.0015)	(0.0013)	(0.0016)	(0.0013)	(0.0013)
ling. fractionalization	0.0442	0.4315**	0.1481	0.0503	0.1518	0.1528
	(0.1730)	(0.1640)	(0.1523)	(0.1419)	(0.1532)	(0.1518)
farm size	0.0925*	0.1956***	0.1408***	0.1289***	0.1414***	0.1406***
	(0.0359)	(0.0367)	(0.0348)	(0.0365)	(0.0347)	(0.0352)
farm size squared	−0.0037**	−0.0057***	−0.0039**	−0.0036**	−0.0039**	−0.0039**
	(0.0013)	(0.0014)	(0.0013)	(0.0013)	(0.0013)	(0.0013)
land inequality	−0.4752*	−0.9721***	−0.8487***	−0.8878***	−0.8632***	−0.8521***
	(0.2391)	(0.2189)	(0.2019)	(0.2008)	(0.2122)	(0.2009)
urbanization			−1.3505***	−1.2294***	−1.3498***	−1.3493***
			(0.2740)	(0.2581)	(0.2741)	(0.2740)
suburbs			−0.1205	−0.0615	−0.1204	−0.1210
			(0.1184)	(0.1109)	(0.1183)	(0.1186)
industrial employment				−1.0096+		
				(0.5608)		
market potential				−0.5397*		
				(0.2288)		
Raiffeisen distance					0.0105	
					(0.0505)	
District F.E.		✓	✓	✓	✓	✓
Year F.E.						✓
Observations	14632	14632	14632	14632	14632	14632
No. of cooperatives	4941	4941	4941	4941	4941	4941

Sample: Six eastern provinces of Prussia, 1852–1913. Dep. Var. year-on-year change in number of credit cooperatives per county; grain prices refers to the year-on-year change of the moving average of weighted rye and wheat prices lagged over the past five years, deflated by the consumer price index. Limited liability, central bank and incorporation are time varying policy dummies. Other variables are time invariant. See tables B1 and B.2 in the online appendix for further definitions. Poisson model with standard errors clustered at county level (236 counties). Fixed effects at the level of 14 districts. All regressions exclude Berlin. Standard errors in parentheses: + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

ties, and because wheat and rye prices may move differently.[24] We therefore weigh relative wheat and rye prices in our price index using the share of total land in each county dedicated to wheat and rye production $Z_i^{rye,\,wheat}$. This captures the degree to which each county was potentially affected by global grain price movements:

$$P_{i,t}^{rel.\,grain} = \left(\frac{P_t^{wheat}}{P_t^{CPI}} Z_i^{wheat}\right) + \left(\frac{P_t^{rye}}{P_t^{CPI}} Z_i^{rye}\right) \tag{8}$$

In most specifications, we employ land use data from 1878, which is the earliest year for which county-level data is available from Meitzen (1901). As land use itself could change as a response to price shocks, we also calculate land use at the beginning of our sample (1852) from province level data, but this does not affect the estimates (see column (2)–(4) in table F3).[25]

The resulting price index was plotted in Fig. 1c in the introduction. Clearly, the series is non-stationary, with the attendant risk of spurious results. We therefore difference the series once to arrive at $\Delta P_{i,t}^{rel.\,grain}$. Finally, it seems reasonable to expect that not just the contemporaneous change in prices affects credit cooperative growth, but recent lags as well. This may be because farmers react to a number of adverse price shocks over recent years, or because setting up a cooperative takes time. Our main explanatory variable "grain prices" is therefore a moving total of $\Delta P_{i,t}^{rel.\,grain}$ over the current and past five years.[26] The differenced price series, and its moving five-year total are plotted in Figure C2.

3.3.2. Asset size and inequality

Our conceptual framework stresses the role of mean assets per capita q in a region, as well as asset inequality. We use farm size as a proxy for assets. There are three reasons for doing so. Firstly, in an agricultural economy, land is arguably the most important asset. Other assets, such as farm buildings are likely to be correlated with land holdings (Eddie, 2008). Secondly, land is used as collateral in financial transactions, for

[24] Map E.1c displays the cross-sectional variation in grain cultivation at the county level. The majority of this was rye cultivation, although there are large differences between counties in the crop composition (see summary statistics in table B2). Graph C.1 displays the movement of the difference between rye and wheat prices over time.

[25] The results also carry through when we instrument land use with exogenous variation in soil suitability in column (5) in table F3. Specifically, we instrument Z_i^{wheat} with the share of fertile land from Meitzen (1901), and Z_i^{rye} with the remaining share of land.

[26] The results are very similar if we enter each lag of $\Delta P_{i,t}^{rel.\,grain}$ separately into the regression, but the exposition and reading of the results is significantly eased by using one cumulative total. The results are also robust to including a different number of lags than the 5-period cut-off.

Table 2
Determinants of credit cooperative growth, county level, Prussia: **Land inequality**.

	(1) Top two land inequality	(2) Gini land inequality	(3) Gini with decade effects	(4) Instrumenting land inequality	(5) Instrument & soil quality	(6) Instrument & public goods	(7) Instrument & state funding
	Dependent Variable: Number of new credit cooperatives per county						
Δ grain prices	−38.4079***	−38.4126***	−29.4269***	−38.3059***	−38.3129***	−38.2945***	−38.5345***
	(3.0645)	(3.0714)	(2.9270)	(3.0319)	(3.0340)	(3.0357)	(3.0323)
population	0.0067***	0.0068***	0.0068***	0.0062***	0.0062***	0.0063***	0.0045**
	(0.0013)	(0.0014)	(0.0014)	(0.0015)	(0.0015)	(0.0015)	(0.0015)
ling. fractionalization	0.0926	0.0781	0.0826	0.4646**	0.4850**	0.3864*	0.3238*
	(0.1548)	(0.1536)	(0.1537)	(0.1793)	(0.1762)	(0.1698)	(0.1571)
farm size	0.1361***	0.1253***	0.1276***	0.2094***	0.2034***	0.2075***	0.1072**
	(0.0346)	(0.0342)	(0.0342)	(0.0451)	(0.0431)	(0.0434)	(0.0355)
farm size squared	−0.0037**	−0.0036**	−0.0037**	−0.0059***	−0.0058***	−0.0058***	−0.0032**
	(0.0012)	(0.0012)	(0.0012)	(0.0015)	(0.0015)	(0.0014)	(0.0011)
land inequality	−0.7688**			−1.3200*	−1.1674*	−1.2606*	−0.9868*
	(0.2348)			(0.6496)	(0.5729)	(0.6046)	(0.4948)
Gini		−1.2476*					
		(0.4960)					
−1854 × Gini			−7.8558***				
			(1.7538)				
1855-1864 × Gini			−3.3025**				
			(1.1036)				
1865-1874 × Gini			−3.0109**				
			(1.0923)				
1875-1884 × Gini			−4.3488***				
			(1.1016)				
1885-1894 × Gini			−3.9138***				
			(0.8789)				
1895-1904 × Gini			−0.5905				
			(0.5375)				
1905-1913 × Gini			−1.1339*				
			(0.5480)				
Berlin distance				0.0504			
				(0.0694)			
soil quality					0.1195		
					(0.1513)		
pupil-teacher ratio						0.0055	
						(0.0042)	
state public spending							0.1441
							(0.1760)
District F.E.	✓	✓	✓	✓	✓	✓	✓
Urbanization controls	✓	✓	✓				✓
Policy dummies	✓	✓	✓	✓	✓	✓	✓
Observations	14632	14632	14632	14632	14632	14632	14632
No. of cooperatives	4941	4941	4941	4941	4941	4941	4941

Sample: Six eastern provinces of Prussia, 1852–1913. Dep. Var. year-on-year change in number of credit cooperatives per county; grain prices refers to the year-on-year change of the moving average of weighted rye and wheat prices lagged over the past five years, deflated by the consumer price index. Policy dummies refers to time varying dummies capturing reform of limited liability, central bank and incorporation law. Other variables are time invariant. Urbanization controls include the share of the population resident in urban areas and a dummy for suburban counties. Specifications (4)–(7) instrument land inequality (share of large estates) with distance to medieval military settlements. See tables B1 and B.2 in the online appendix for further definitions. (1), (2) & (3) Poisson, (4)–(7) GMM. Standard errors clustered at county level (236 counties). Fixed effects at the level of 14 districts. All regressions exclude Berlin. Standard errors in parentheses: + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

example when taking a bank loan. Thirdly, land holdings are likely to be correlated with income, which determines the ability to repay loans and make savings deposits.[27] We utilize the 1882 land census as collected in the iPEH-Database (Becker et al., 2014) to calculate mean farm size by simply dividing a county's total farmland in hectares by the total number of farms.

For asset inequality, we utilize the same data source. The data divides all farmland in a county into six size categories, and provides the amount of land in each category. Our preferred measure is simply the share of arable land in a county held by estates in the largest category (over 100 ha). This is consistent with the spirit of $\overline{q}$ in the model and the treatment of inequality as upper-tail inequality.[28] As a robustness check, we also compute a Gini index of farm size inequality, by

[27] Losch et al. (2011) report that farm size is still the major determinant of rural incomes in a broad cross section of developing countries. In our context, one might worry that farm value, rather than size, matters. The value of land could trend downward as demand for land slackens. However, interacting farm sizes with decade time trends does not reveal such patterns. Similarly, proxying for farm values by interacting farm size with soil quality, industrialization, demographic variables, or market potential (Kopsidis and Wolf, 2012) does not greatly affect the coefficients on farm size (table F2).

[28] Land in this category was officially denoted as *Grossgrundbesitz* (large estate) in Prussia. The measure largely follows Cinirella and Hornung (2016), although the authors use the share of farms rather than the share of land. Note that the threshold for classifying as a large estate in Prussia is quite distinct, and is likely to be much larger, than the collateral threshold required for bank loans q_T in the model.

Table 3
Determinants of credit cooperative growth, county level, Prussia: **Prices and other time varying variables**.

	(1) Prices, county fixed effects	(2) Controlling for bank interest rates	(3) Controlling for GDP growth	(4) Instrumenting with US prices	(5) Instrumenting with UK interest rates	(6) Instrumenting with UK growth rates	(7) Controlling for price volatility
	Dependent Variable: Number of new credit cooperatives per county						
Δ grain prices	−37.21***	−35.30***	−37.99***	−46.74***	−33.96***	−33.35***	−32.30***
	(3.16)	(3.38)	(3.02)	(4.87)	(3.33)	(3.54)	(2.86)
limited liability	1.58***	1.87***	1.59***	1.65***	1.85***	1.77***	1.59***
	(0.15)	(0.22)	(0.15)	(0.15)	(0.21)	(0.26)	(0.15)
central bank	1.39***	1.63***	1.39***	1.35***	1.68***	1.42***	1.48***
	(0.14)	(0.14)	(0.14)	(0.14)	(0.14)	(0.14)	(0.14)
incorporation	−0.12		0.06	−0.03		0.028	0.32*
	(0.13)		(0.13)	(0.14)		(0.14)	(0.15)
interest rates		−0.28***			−0.34***		
		(0.033)			(0.040)		
GDP growth			0.0112			0.3404*	
			(0.0133)			(0.1410)	
population				0.0067***	0.0064***	0.0067***	
				(0.0013)	(0.0014)	(0.0013)	
ling. frac.				0.1268	0.2043	0.1332	
				(0.1515)	(0.1604)	(0.1517)	
farm size				0.1345***	0.1341***	0.1394***	
				(0.0339)	(0.0348)	(0.0340)	
farm size sq.				−0.0037**	−0.0035**	−0.0038**	
				(0.0012)	(0.0013)	(0.0012)	
land inequality				−0.8173***	−0.9006***	−0.8145***	
				(0.2013)	(0.2121)	(0.2017)	
price volatility							8.07***
							(1.66)
District F.E.				✓	✓	✓	
County F.E.	✓	✓	✓				✓
Urbanization controls				✓	✓	✓	
Observations	13572	8740	14508	13572	8740	14508	14508
No. of cooperatives	4940	4591	4941	4940	4714	4591	4941

Sample: Six eastern provinces of Prussia, 1852–1913. Dep. Var.: year-on-year change in number of credit cooperatives per county; grain prices refers to the year-on-year change of the moving average of weighted rye and wheat prices lagged over the past five years, deflated by the consumer price index. Price volatility is the five year standard deviation of the change in deflated rye and week prices. Limited liability, central bank and incorporation are time varying policy dummies. GDP growth is annual change in German GDP, interest rate is the lending rate of German commercial banks (available from 1876). Other variables are time invariant. Urbanization controls include the share of the population resident in urban areas and a dummy for suburban counties. See tables B1 and B.2 in the online appendix for further definitions. Specification (3) instruments relative grain prices using US prices (available from 1856), (4) instruments interest rates using the UK Bank of England discount rate, (5) instruments GDP growth with UK GDP growth lagged by 1 year. Specifications (1), (2), (3), (7) are county fixed effects Poisson regressions (234 counties). (4), (5) and (6) are GMM with fixed effects at the level of 14 districts. Standard errors clustered at county level (236 counties). All regressions exclude Berlin. Standard errors in parentheses: + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

calculating the area under a discrete Lorenz curve of all six size categories, as well as experimenting with different cut-offs, but the results turn out to be insensitive to these alterations.

3.3.3. Population and heterogeneity

All demographic data stems from the official Prussian censuses, collected by Belzyt (1998). The censuses provide population totals, as well as the breakdown by ethnic and religious groups for each county. For most counties, ten census results are available between 1858 and 1910. We linearly interpolate the data between census years.[29] Although we sometimes allow population to vary over time, we use fixed 1890 population figures in most specifications. This is the census that is available for all counties, and avoids spurious results due to trending population numbers. We also use 1852 numbers (column (6) in table F1) as a robustness check to forestall the possibility that population itself might adjust to price shocks, for example through migration.

According to our theory, communication costs are more likely to be present when both ethnicities are of roughly equal size. An index of ethnic fractionalization is one way of capturing these non-linearities. We calculate a generalized formulation of the heterogeneity index in equation (3) that allows for more than two groups (although in the majority of counties, Germans and Poles are the only groups present). For county i we compute our fractionalization index as:

$$frac_i = 1 - \sum_{k=1}^{K} \left(\frac{N_{k,i}}{N_i} \right)^2 \tag{9}$$

where k is a linguistic or religious group. An alternative is a polarization index, as proposed by Esteban et al. (2012), which may be more relevant when focusing on antagonism between two groups, such as Germans and Poles.[30] However, Esteban et al. (2012) emphasize that polarization is a predictor of conflict over public, rather than private goods, and it is the latter that we are primarily interested in. We therefore employ the fractionalization index in most applications.

3.3.4. Banks and other controls

In our setting, controlling for bank presence is important because savings banks were the main substitute to co-ops as sources of finance

[29] We use the territorial definitions of counties in 1900. In cases where county boundaries were redrawn, we extrapolate the population shares backwards using the constituent territories' growth rates, with the relative population shares as weights. This yields stable units of observation over time.

[30] See Cinirella and Schueler (2016), who make this point in a Prussian context specifically.

19th Century Prussia. In line with our conceptual framework, however, they did not generally target individual small farmers either as savers or borrowers (Burhop, 2011; Born, 1985). Moreover, as they could supply credit to cooperatives, they sometimes acted as complementary institutions. We use the founding date of all savings banks present in the year 1909 (Evert, 1911) to construct a panel of these institutions.

Most other county-level control variables are extracted from the aforementioned iPEH-Database (Becker et al., 2014). A small number of variables (urbanization and soil quality) are taken from Meitzen (1901). Tables B1 and B2 provide summary statistics and a short description of all variables. The main variables of interest are also mapped in Appendix A.

3.4. Empirical strategy

3.4.1. Benchmark specification

Based on the comparative statics properties in PROPOSITION 1, we know that a change in grain prices should determine market entry by cooperatives, conditional on fixed county characteristics. Our empirical benchmark specification explains the formation of credit cooperatives at the county level from 1852, the year of the first foundation, to 1913, the last year before World War I. Formally:

$$\Delta C_{i(j),\,t} = \beta_0 + \beta_1 \Delta P_{i,\,t}^{rel.\,grain} + \beta_2 N_i + \beta_3 H_i + \beta_4 q + \beta_5 q^2 + \beta_6 \overline{q} + V_t + W_i + \alpha_j + \epsilon_{i,\,t} \quad (10)$$

where $\Delta C_{i(j),\,t}$ is the growth in the number of credit cooperatives in county i in year t. The main explanatory variables are those developed in our conceptual framework: β_1 captures the role of grain price changes, β_2 and β_3 examine the effect of population and population heterogeneity.[31] β_4 investigates the effect of mean assets, measured by mean farm size, and β_5 takes account of the hypothesized hump-shaped relationship between mean assets and cooperatives. Finally, β_6 tests the association between asset inequality $\overline{q}$ and co-op growth. In the benchmark specification, we include two further control variables in W_i: urbanization (the share of a county's population living in urban settlements), and a dummy for "surburban" counties, which we define as those that encompass a town not part of that county. In some specifications, we control for the number of banks per county at the start of the panel.[32]

We always include three time varying dummies in V_t that capture important policy changes affecting the legal environment in which cooperatives could operate (Faust, 1965). These are the 1867 incorporation law bestowing legal personhood unto cooperatives, the 1889 law allowing co-ops to elect limited liability on members' assets, and the 1895 founding of the *Preussenkasse*, a central public bank set up to extend loans to cooperatives experiencing short term liquidity crises.[33] In some specifications, we also test for the other time-varying factor from our model: annual bank interest rates r_b from Bundesbank (1976). We also test for the effect of GDP growth, which has been shown to affect the growth of MFIs (Ahlin et al., 2011).

The dependent variable in equation (10) is a discrete non-negative count, and therefore estimated using a Poisson (Cameron and Trivedi, 2015). In most specifications, we pool data across counties and years, although we always employ fixed effects α_j at level of the 14 districts. This is the administrative layer between our 6 provinces and 236 counties, and groups together historically similar counties. We also estimate some models with county or year fixed effects. Standard errors are clustered at the county level.

3.4.2. Causal identification

To what extent will a regression of credit cooperatives on grain prices yield causal estimates? Our counties are too small to have influenced the German price level, so that we can rule out reverse causality. However, third factors, such as large regional harvest shocks, could have affected aggregate grain supply and hence prices. Such harvest shocks could also directly drive credit demand. In this case β_1 in equation (10) may not be picking up the causal effect of prices on credit cooperatives.[34] Ultimately, we are interested in falling grain prices as a feature of structural transformation. These price changes were the result of evolving domestic demand patterns and international integration. We therefore make use of a theoretical result from the literature on structural transformation derived by Teignier (2018): In a small, land scarce, and reasonably open economy, domestic relative prices should be determined by (lower) world market prices for agricultural goods. The resulting lower price ratio can drive structural change. As a grain producer Prussia, and Germany in its entirety, were small relative to the land abundant producers in Argentina, Russia, and the USA that were increasingly dominating German grain markets. Moreover, as non-perishable staple grains were more easily tradeable than dairy or meat, import competition was particularly severe for these goods (Findlay and O'Rourke, 2009). We therefore use US prices for wheat and rye (Jacks, 2013) as an instrument for German prices. The US was one of the largest producers of staple grains in the world, which ensures the instrument is relevant.[35] Moreover, the exclusion restriction that US grain prices affect the creation of local credit cooperatives only through its effect on German prices is likely to hold given that US and local Prussian supply conditions are unlikely to be correlated.

We also exploit the international environment of 19th century globalization to identify the causal effect of other macroeconomic variables. We instrument German bank interest rates, which could conceivably be driven by competition from expanding credit cooperatives, by the Bank of England's official discount rate. The Bank of England acted as the "conductor" of the international monetary system of the 19th century, and national central banks often followed its rate setting policy, transmitting its rates to domestic markets (Eichengreen, 1987). Figure C5 shows the strong comovement between the Bank of England and German commercial rates.

Similarly, as discussed by Ahlin et al. (2011), a regression of MFI outcomes on GDP growth may not produce causal effects due to omitted variables. In our setting, one possibility may be that, as rural incomes are increasing, this increases GDP growth while at the same time decreasing rural credit demand.[36] We therefore instrument German GDP growth with UK GDP growth, lagged one period. The UK was Europe's largest economy at the time, and Germany's principal trade partner, making the instrument relevant. It also seems unlikely that UK incomes of the past year were driven by current German demand conditions.

[31] Clearly, a higher rate of co-op formation is mechanically expected in larger counties, and β_2 should not be interpreted as measuring the impact of group size N_c from the model.

[32] Although variables like population or banks could be allowed to vary over time, we prefer keep them fixed in most specifications to mitigate their potential endogeneity to grain price shocks, which would lead to a "bad controls" problem.

[33] See Guinnane (2004) for a discussion of the actual role of the *Preussenkasse* and the middle-tier regional "central" banks. Econometrically, these policy dummies account for the higher mean growth rate of co-ops after 1890 visible in Fig. 1d.

[34] Specifically, if domestic grain supply is negatively related to gain prices and demand for credit, β_1 would understate the true negative effect of grain prices on credit cooperatives.

[35] Figure C3 shows the degree to which US and German prices are indeed correlated. Note that what matters for our instrument is not whether US and German price levels were the same (given trade costs and tariffs), but that the changes were closely correlated.

[36] This would imply that the coefficient would be biased towards 0, which is what we in fact observe in Table 3 (columns (3) and (6)).

A different strategy is required for cross-sectional variables, such as land inequality. Here we exploit the unique history of settlement of the Prussian eastern provinces. Specifically, we focus on the fact that these territories were settled by colonists from the High Middle Ages onwards, and that there were, roughly, two motivations for settlement. One was economic: farmers, craftsmen and merchants were drawn to the East because it offered relatively sparsely populated land. The other was military: parts of the East were populated by pagan peoples, which drew a number of crusading military orders to the region, the most notorious being the Teutonic Order (Herrman, 2015). These crusaders established large landed estates centered around military strongholds on the land conquered from pagan peoples. Although the crusading orders themselves had been abandoned by 1525, and many of their strongholds were ruins by the 19th century, their settlements planted the seed for the unequal land distribution in the region (Torbus, 1998). We therefore measure the distance from each county capital to these historical military strongholds to arrive at an instrument for land inequality.[37] The instrument's validity is predicated on the notion in the historical literature that the location of these military settlements were driven by the shifting security situation of the middle ages, rather than the economic considerations of the 19th century (see Torbus (1998, p. 56–57), Leighton (2016) and Piana and Carlsson (2014)).[38]

For other variables, we are not able to employ instrumental variables. Yet we do utilize another key feature of our historical setting by exploiting the precedence in timing of our covariates. For example, we can hold county-level ethno-linguistic heterogeneity or the development of the banking system fixed at their 1852 values, before credit cooperatives even existed. Although this does not completely rule out the possibility of "deep" county-specific omitted variables, it does decrease their likelihood and rule out reverse causality.[39]

4. Results

4.1. Benchmark results

The main results in Table 1 strongly support the predictions of our theoretical framework. We commence in column (3) by showing the results of the full benchmark specification derived in section 3.4.1. The coefficient on grain prices is negative and strongly statistically significant, suggesting that falling relative grain prices played a role in the rise of credit cooperatives in counties heavily reliant on grain production. Introducing year fixed effects decreases the precision of the estimator, but the coefficient is left intact (column (6)).

The coefficient on farm size is positive while its quadratic term is negative: high levels of rural assets encourage the formation of credit cooperatives, but beyond a certain threshold, asset wealth actually decreases cooperative growth. This is consistent with the hypothesized hump-shaped relationship between private assets and small-scale credit arising from the presence of bank lending with collateral requirements. The coefficient on land inequality is negative, showing that counties with a higher share of land held in large estates provided a less hospitable environment for the growth of microfinance institutions.

The impact of linguistic fractionalization is as ambiguous empirically as it is in theory. Although the coefficient is generally positive, in line with PROPOSITION 2, it is only statistically significant in column (2) before the introduction of urbanization controls. As the coefficient on the urbanization variable shows, most credit cooperatives thrived in rural areas. These rural areas were also disproportionally fractionalized, accounting for the strength of the positive correlation in (2).

As can be expected, population is a strong predictor of cooperative growth, but this relationship is partly mechanical. Finally, the time-varying policy dummies indicate that the institutional environment had a strong effect on co-op growth: the introduction of limited liability laws, and the founding of the cooperative central bank seem to have played a large role in encouraging credit cooperative formation.

The above results are robust to the introduction of a wider set of control variables, for example the percentage of the workforce employed in manufacturing, and market potential in column (4).[40] The variables themselves emphasize the idea that, conditional on the distribution of private land assets, credit cooperatives were more likely to form in less industrialized and more remote areas. In column (5) we also examine whether the geographical spread of the cooperative idea itself mattered (Jensen et al., 2018), but find that credit cooperatives were no more likely to form if they were closer to Neuwied, the location where Raiffeisen propagated his ideas.

We include more controls, including 1892 wage levels, labour income growth between 1880 and 1905, migration rates between 1895 and 1905, illiteracy, and population density in table F1, but do not include these in the main specifications as they may be affected by the price shocks themselves. These variables are not time varying, but we get similar results when allowing population to vary over time in column (5) of the same table. Column (6) in table F1 fixes population and fractionalization at their 1852 values. This precludes simultaneity bias, although it may also increase measurement error. Nonetheless, the conclusions of the benchmark regression carry through despite these alterations.

We will now investigate the mechanism behind each of our main variables, and investigate their causal effects.

4.2. Inequality

The benchmark regression uses the share of land in large estates as a measure for land inequality, reflecting the focus in our theoretical framework on the share of top assets $\bar{q}$. In column (1) of Table 2, we extend this definition to include the top two land size categories, with little difference to the results. In column (2) we depart from this approach by taking the whole land distribution of a county into account through a Gini-coefficient. The results are statistically slightly weaker, but still indicate a negative effect of land inequality on rural financialization at conventional significance levels.[41]

However, an important concern could be that our measures of land inequality pick up omitted variables correlated with both land holdings and rural credit. For example, county-level natural endowments, including soil quality, could conceivably affect both measures. As a first step, we check whether our measure behaves in line with the historical evidence on land inequality (Eddie, 2008). We know from existing work

[37] We use a historical atlas (Zeissig, 1964) to code the location of the principal strongholds. The atlas identifies 17 such settlements.

[38] We include soil quality to control for potential economic determinants of settlement, but this does not affect our estimates. Soil quality is used as an instrument for land inequality by Cinirella and Hornung (2016) in their study of education. We do not use it for this purpose, as soil quality will have affected rural cooperatives through influencing the local crop mix (see section 3.3.1.

[39] One important variable that we only observe mid-way in our sample period (1886) is farm size q. However, we also run our benchmark regression solely on data after that year, and find that the coefficient on q has not changed significantly compared to the benchmark, thus suggesting that endogeneity of farm sizes to cooperative foundation is not a major confound (see column (4), table F6).

[40] Market potential is the distance from each county capital to all other German counties and foreign capitals, weighted by the size (population) of these markets (Kopsidis and Wolf, 2012).

[41] We also excluded land inequality or farm size squared from the regression, but found that it left the main effect of farm sizes qualitatively similar. Reversely, we experimented with dropping all land-related variables and instead included all bins of land size save one in the regression. This too produced the predicted hump-shaped relationship between asset size categories and cooperative foundation.

Table 4
Determinants of cooperative growth, county level, Prussia: **Savings banks and other cooperatives**.

	(1) Explaining production cooperatives	(2) production co-ops with fixed credit co-ops	(3) Credit cooperatives with banks	(4) Controlling for bank growth	(5) Controlling for lagged cooperatives	(6) Placebo: other cooperatives
Dependent Variable:	New production co-ops per county		New credit cooperatives per county			Other co-ops
Δ credit cooperatives	0.2186***					
	(0.0197)					
credit cooperatives 1873		0.0981*				
		(0.0476)				
lag credit cooperatives					−0.0119***	
					(0.0034)	
Δ grain prices			−38.4872***	−37.1641***	−34.0767***	6.7631+
			(3.0761)	(3.1637)	(3.1267)	(3.6497)
limited liability			1.5923***	1.5840***	1.5759***	2.7979***
			(0.1539)	(0.1534)	(0.1532)	(0.2047)
central bank			1.3887***	1.4876***	1.5582***	1.2151***
			(0.1411)	(0.1403)	(0.1430)	(0.1130)
incorporation			0.0671	−0.1431	−0.0092	
			(0.1313)	(0.1361)	(0.1322)	
population			0.0077***	0.0076***	0.0068***	0.0053**
			(0.0016)	(0.0015)	(0.0014)	(0.0019)
ling. fractionalization			0.1822	0.0988	0.3132+	−0.5203
			(0.1579)	(0.1640)	(0.1818)	(0.3702)
farm size			0.1353***	0.1402***	0.1568***	0.3012***
			(0.0338)	(0.0359)	(0.0383)	(0.0677)
farm size squared			−0.0037**	−0.0038**	−0.0043**	−0.0078***
			(0.0012)	(0.0013)	(0.0014)	(0.0021)
land inequality			−0.7830***	−0.8917***	−0.9641***	−0.6886
			(0.2049)	(0.2161)	(0.2619)	(0.6314)
banks, 1852			0.1379**	0.1524**	0.1773***	0.1103
			(0.0477)	(0.0504)	(0.0504)	(0.1046)
Δ banks				−0.0065		
				(0.0562)		
income growth			−0.0010	−0.0013		
			(0.0020)	(0.0021)		
market potential			−0.4298*	−0.5100**		
			(0.2003)	(0.1847)		
soil quality			0.0193	0.0573		
			(0.1372)	(0.1433)		
population, 1870		0.0001+				
		(0.0000)				
County F.E.	✓					
District F.E.		✓	✓	✓	✓	✓
Urbanization controls		✓	✓	✓	✓	✓
Observations	11781	9912	14632	12508	14396	14632
No. of cooperatives	898	898	4941	4375	4940	2000

Sample: Six eastern provinces of Prussia, 1873–1913 in (1)–(2); 1852–1913 in (3), (5) & (6), 1852–1909 in (4). Dep. Var. for specifications (1)–(2): year-on-year change in number of production cooperatives per county; (3), (4) and (5): year-on-year change in number of credit cooperatives per county, (6): year-on-year change in number of other cooperatives per county. Banks refers to the time-varying number of banks per county, available before 1909 only, except banks 1852, which refers to the count of banks at that date. See tables B1 and B.2 in the online appendix for further definitions. All specifications are Poisson. Standard errors clustered at county level (236 counties). Fixed effects at the level of 14 districts. All regressions exclude Berlin. Standard errors in parentheses: + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

that even though land inequality itself was not necessarily decreasing during the 19th century, its effect diminished as peasants were gradually freed from their manorial obligations and acquired full ownership rights over plots (Cinirella and Hornung, 2016). Our findings are in line with this interpretation: Once we interact land inequality with a decadal time trend in column (3), we find that the effect of inequality on credit cooperatives does indeed diminish over time, although we cannot rule out attenuation bias here.

Although evidence from time trends makes it less likely that our findings are confounded by county characteristics, the effect of such confounders could conceivably be time-varying too. We therefore utilize distance to medieval military settlements as an instrument for the share of land under large estates as described in section 3.4.2. Column (4) demonstrates that after instrumentation, the coefficient on inequality remains significant and negative.[42] As the location of these settlements was dictated by the military situation facing medieval crusaders, rather than the agricultural conditions prevailing six centuries later, the estimate should now be free from omitted confounders.[43] Moreover,

[42] Similarly, the reduced form relationship (columns (6) and (7) in table F4) shows that distance to medieval settlements predicts co-op formation six centuries later. Note that we do not include urbanization controls in most of the instrumental variable regressions concerning inequality, as patterns of urbanization will have been historically determined by the instrument itself.

[43] One common problem with geographical instruments is that the distance to a single point could be correlated with unobserved spatial variables. In this case, however, we use the sum of distances to 17 settlements, making such coincidental correlation unlikely as long as the initial allocation of settlements is as good as random. Consequently, our instrument is robust to the inclusion of other distance-based measures.

column (5) shows that the instrument is not affected by controlling for soil quality directly.

Although the instrument strengthens the argument that land inequality affects credit cooperatives, it does not pin down a mechanism. While the model emphasizes the impact of land inequality through the distribution of bankable assets, other explanations are plausible. Large landowners could have underprovided public goods such as education, thus hampering co-op growth. They could also have used their political power to stymie the growth of new credit institutions (Rajan and Ramcharan, 2011). Although it is difficult to conclusively rule out these mechanisms, we are able to provide suggestive evidence against them. In column (6) we control for the pupil-teacher ratio as a proxy for educational provision. In column (7) we control for the share of school funding that comes from the state, rather than other (local) sources. This proxies the degree to which large landowners may have substituted for state organs, and thus gauges their political power. In both cases, we find the coefficient on land inequality unaffected.

4.3. Rural transformation

According to our conceptual framework, a decline in staple prices encourages farmers to found microfinance institutions. These institutions allow credit-constrained producers to enter the production of capital-intensive agricultural goods, driving rural transformation. In this section, we show that the empirical evidence strongly supports the causal impact of relative grain prices on the formation of credit cooperatives. There is also suggestive evidence that these cooperatives aided rural transformation.

We first examine the effect of relative grain price changes in a setting with county fixed effects in column (1) of Table 3. The coefficient on grain prices is virtually unchanged from the benchmark regression.[44] However, German grain prices could be endogenous to domestic supply conditions. Using US grain prices as an instrument in (4) takes care of this concern. Movements in US prices impacted German grain price changes, and this encouraged farmers in those counties of Prussia heavily reliant on grain production to set up credit cooperatives. The coefficient on grain prices increases, suggesting that domestic harvest shocks affecting prices and credit demand simultaneously biased the coefficient towards zero.[45] We obtain similar conclusions in the reduced form (column (2) in table F4), as well as when instrumenting rye and wheat cultivation with exogenous variation in soil quality (column (5) in table F3). Theoretically, trade policy could have affected the pace of financial development in rural areas by shielding domestic grain producers. In column (5) and (6) of table F6 we separately consider the years before and after Germany's landmark 1879 tariffs (Lehmann, 2010). Yet we find the direct effect of US grain prices to be *higher* in the decades after the tariffs, suggesting technology-driven market integration outpaced policy markers.

In the model, bank lending rates act as a floor for cooperative lending rates. High bank interest rates will therefore negatively impact the formation of credit unions. This assertion finds support in column (2) in Table 3, where the annual mean interest rates of German commercial banks is shown to be negatively correlated with the co-op growth rate. Exploiting exogenous interest rate shocks from the Bank of England, the results in column (5) suggest that this relationship is causal: increases in the discount rate in the UK changed the lending practices of German banks, which decreased the attractiveness of setting up credit cooperatives.

Although not explicitly a part of our model, we also investigate the role of the GDP growth rate on microcredit, as rising incomes are intimately related to structural transformation (column (3)). In column (6), we instrument German GDP growth with lagged UK GDP growth, and find there to be a statistically significant impact on co-op formation. This supports the conclusions Ahlin et al. (2011) draw from correlational evidence on modern data. We interpret this result as being related to structural transformation: an exogenous increase in German per capita incomes shifts relative demand away from staples. With prices fixed by world markets, this implies a smaller output share for staples, increasing the demand for loans to aid exiting that sector. This is supported by the observation that the coefficient on GDP growth is not different from 0 if we ignore endogeneity in column (3): domestic German GDP growth partially reflects increasing farm incomes, and this decreases credit demand.

Although the effect of grain prices on credit cooperatives is robust, there are alternative interpretations of this relationship that do not draw on structural change. Rather than exiting grain production, farmers could have accessed credit to smooth consumption or insure against risk (Fafchamps and Gubert, 2007). Although cooperatives did mainly extend loans for productive purposes rather than for consumption or debt refinancing, this rule was not always followed (Guinnane, 2001). We therefore include grain price volatility as a covariate in column (7) in Table 3.[46] The coefficient on price volatility has a positive sign and is statistically significant, suggesting that price volatility is an alternative channel leading to the development of rural credit markets. However, the results for grain prices remain largely unaffected. Even more, the effect of price volatility is economically small (section 4.6), and price volatility was generally declining as market integration proceeded (figure C4).

In Table 4 we provide additional indirect evidence on the role of credit cooperatives in helping farmers to exit staple production. Because annual data on changes in agricultural output at the county level is lacking, we use the founding of production cooperatives as a proxy for annual changes in rural production. These cooperatives produced income elastic goods such as dairy and meat. Column (1) shows that a higher growth rate of credit cooperatives is indeed associated with a contemporaneous increase in the foundation of production cooperatives. Although this regression controls for county-fixed effects, it could still be that time varying omitted factors drive both growth rates. In column (2), we rule this out by fixing the number of credit cooperatives at the 1873 value, the year in which the first production cooperative was founded (see Fig. 1d).

In Table 5, we turn to direct evidence for rural transformation by regressing the change in livestock numbers between 1868 and 1906 on the growth in credit cooperatives during this period at the county level. The correlation is statistically strong and economically meaningful for all three livestock categories: dairy cattle, beef cattle, and pigs.[47] One additional cooperative is associated with 130 additional heads of dairy cattle during this period, and with more than 400 pigs. The results are robust to controlling for a range of covariates, including 1868 livestock levels, as well as urbanization levels and income and population growth. Although these particular results are suggestive rather than causal, they do lend support to the interpretation that credit cooperatives helped grain producing regions

[44] We use the Poisson conditional fixed effect model with the robust sandwich error estimator proposed by Hausman et al. (1984), which has the benefit of being robust to a misspecification of the underlying count distribution, serial correlation, as well as time-invariant spatial dependence (Bertanha and Moser, 2016). We loose observations from two counties with zero co-op growth, and from four years without US price data. We obtain similar results in a specification allowing for both county and year fixed effects (column (1) in table F3).

[45] Incidental parameters imply that we cannot use county fixed effects in this GMM set-up, so that the IV results are comparable to the benchmark in column (3), Table 1 rather than the fixed effects specification.

[46] The variable is calculated as the standard deviation of changes in our relative grain price index over the past five years.

[47] Data on livestock numbers stems from the Prussian cattle censuses (Meitzen, 1901; Preußisches Statistisches Landesamt, 1908). We calculate beef cattle as total cattle subtracting cows and calves.

Table 5
Non-grain rural production, county level, Prussia: **Change in livestock 1868–1906**.

	(1) Dairy cattle	(2) Beef cattle	(3) Pigs	(4) Dairy cattle, controls	(5) Beef cattle, controls	(6) Pigs, controls
	Dependent Variable: Δ livestock headcount 1868–1906					
Δ Credit cooperatives 1868–1906	131.14*	297.51***	474.8***	131.2**	266.7***	416.9***
	(51.5)	(54.0)	(77.8)	(42.9)	(47.9)	(71.5)
dairy cattle, 1868	−0.013			−0.115		
	(0.113)			(0.132)		
beef cattle, 1868		−0.546**			−0.616***	
		(0.171)			(0.166)	
pigs, 1868			0.991***			0.865***
			(0.187)			(0.196)
population growth				−532.2+	−1217.1*	−689.7
				(274.4)	(495.9)	(463.9)
income growth per capita				−13.25	−34.42*	4.32
				(12.6)	(15.9)	(42.9)
District F.E.	✓	✓	✓	✓	✓	✓
Urbanization controls				✓	✓	✓
Observations	236	236	236	236	236	236
R-squared	0.446	0.426	0.596	0.513	0.482	0.618

Sample: Six eastern provinces of Prussia 1868–1906. Dep. Var. total change in headcount of lifestock (dairy cattle, beef cattle, pigs) per county over the entire period. Urbanization controls include the share of the population resident in urban areas and a dummy for suburban counties; See tables B1 and B.2 for further definitions. All regressions are OLS with fixed effects at the level of 14 districts with robust standard errors. All regressions exclude Berlin. Standard errors in parentheses: + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

to diversify farm output and switch to more capital-intensive production.[48]

How does the availability of other financial institutions impact the development of small-scale credit? In our theoretical framework, banks and microcredit institutions serve segmented markets of individual lenders due to the presence of collateral requirements. For borrowers at the threshold, one would expect banks credit and micro credit to be substitutes (Cull et al., 2014; Colvin and McLaughlin, 2014). However, individuals founded cooperatives partly to access bank loans as a group, in which case both institutions would exercise complementary roles (Ahlin et al., 2011; Périlleux et al., 2016). If the latter is true, one might also be concerned that our count of cooperatives simply captures regional financial development. The results in column (3) of Table 4 show that this concern is unsubstantiated: controlling for the number of banks at the county level at the start of the sample does not affect the other coefficients. The coefficient on the number of banks does support the interpretation of complementarity, although it cannot be ruled out that omitted variables at the county level affect the creation of both institutions. Similarly, controlling for growth in the number of banks in column (4) does not change these conclusions, and the results continue to show that small scale credit was still more likely to develop in counties where banks were present early on.

The market entry of a credit cooperative may also be conditional on the existence of other credit cooperatives. The presence of existing cooperatives could encourage the formation of new cooperatives through learning and demonstration effects, or make market entry less attractive through competition and market saturation effects. To this end, column (5) includes the stock of cooperatives, lagged one period, as a regressor. The coefficient is negative and statistically significant, suggesting the presence of saturation effects. The other coefficients are largely unchanged.[49]

Finally, we might be concerned that our results are driven by factors that determine cooperative activity over time more generally, but that have nothing to do with credit cooperatives per se. This we investigate with a placebo test in column (6), where we regress the change in the number of all non-credit cooperatives on our benchmark variables.[50] Grain prices are, if anything, positively related to the growth of these other cooperatives, which may reflect income effects. Banks are now no longer statistically significant.

4.4. Ethno-linguistic heterogeneity

Our conceptual framework showed that heterogeneity among the population could hinder the appearance of rural small scale credit by imposing high costs on transactions between group members (equation (6)). On the other hand, per PROPOSITION 2, these costs can also provide an incentive to create a larger number of ethnically segregated cooperatives. There is indeed historiographical evidence that credit cooperatives in Prussia's eastern provinces became heavily segregated in the last decades of the 19th century (Lorenz and Müller, 2006). However, the benchmark results did not support a strong role for ethno-linguistic fractionalization in affecting cooperative growth. Column (1) of Table 6 shows that the same is true for religious fractionalization. Although the coefficient now has a negative sign, it is still not statistically significant. This does not change when we use polarization measures in column (2). We cannot attribute these results to the endogeneity of the population distribution to price shocks or cooperative growth itself (column (6), table F1).

[48] We find, on the contrary, little evidence that credit cooperatives are associated with stalling structural transformation in the wider economy, for example by hindering emigration or industrial employment in counties where they were active.

[49] Splitting the sample at the spike in the year 1894 reveals that the saturation effect is only present in later years, whereas the coefficient on lagged stocks in earlier years is positive and thus consistent with demonstration effects. We also included the lagged *growth* of cooperatives as regressors in an AR(5) and AR(10) model, but this similarly left the coefficients on prices and the cross-sectional variables of interest unaltered.

[50] Other cooperatives include production, consumption, and resource cooperatives. The latter two were mainly concerned with attaining monopsony power in purchasing consumption goods or raw materials.

Table 6
Determinants of credit cooperative growth, county level, Prussia: **Ethno-linguistic heterogeneity**.

	(1) Religious fractionalization	(2) Linguistic polarization	(3) Share of Polish speakers	(4) Share of Protestants	(5) Inequality interaction	(6) Ethnic inequality
	Dependent Variable: Number of new credit cooperatives per county					
Δ grain prices	−38.4109***	−38.4040***	−38.3615***	−38.4115***	−38.4364***	−38.4159***
	(3.0575)	(3.0580)	(3.0464)	(3.0583)	(3.0635)	(3.0509)
population	0.0069***	0.0068***	0.0067***	0.0069***	0.0067***	0.0066***
	(0.0013)	(0.0013)	(0.0013)	(0.0013)	(0.0013)	(0.0013)
ling. fractionalization			0.3259+		−0.2814	0.1686
			(0.1682)		(0.4115)	(0.1540)
farm size	0.1422***	0.1403***	0.1327***	0.1422***	0.1361***	0.1387***
	(0.0345)	(0.0347)	(0.0347)	(0.0343)	(0.0345)	(0.0348)
farm size sq	−0.0039**	−0.0039**	−0.0037**	−0.0039**	−0.0037**	−0.0038**
	(0.0012)	(0.0013)	(0.0012)	(0.0012)	(0.0013)	(0.0013)
land inequality	−0.8495***	−0.8465***	−0.7204***	−0.8511***	−1.0044***	−0.8324***
	(0.2003)	(0.2019)	(0.2163)	(0.2005)	(0.2809)	(0.2045)
rel. fractionalization	−0.2082			−0.2106		
	(0.2789)			(0.2740)		
ling. polarization		0.0746				
		(0.0805)				
Polish share			−0.2996*			
			(0.1345)			
Protestant share				0.0049		
				(0.1623)		
ling. frac. × land inequality					1.0399	
					(0.9815)	
religious inequality						0.4579
						(0.4674)
District F.E.	✓	✓	✓	✓	✓	✓
Urbanization controls	✓	✓	✓	✓	✓	✓
Policy dummies	✓	✓	✓	✓	✓	✓
Observations	14632	14632	14632	14632	14632	14632
No. of cooperatives	4941	4941	4941	4941	4941	4941

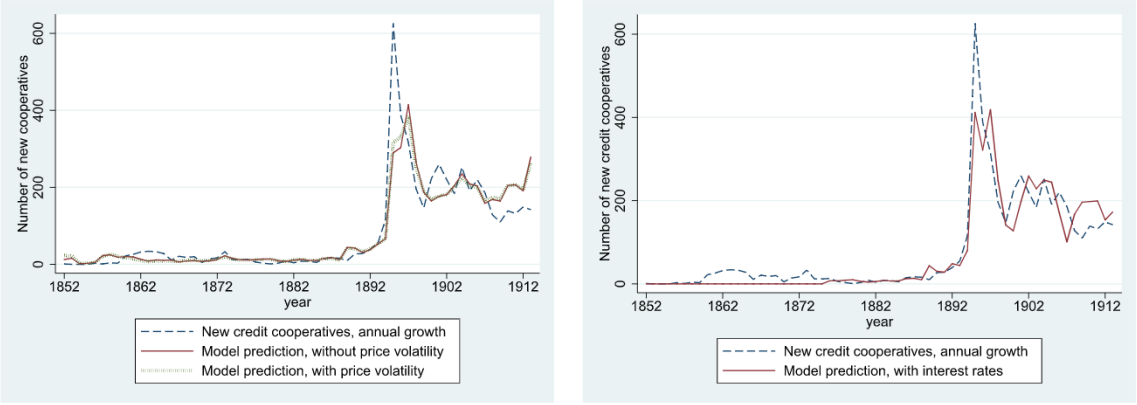
Sample: Six eastern provinces of Prussia, 1852–1913. Dep. Var. year-on-year change in number of credit cooperatives per county; grain prices refers to the year-on-year change of the moving average of weighted rye and wheat prices lagged over the past five years, deflated by the consumer price index. Policy dummies refers to time varying dummies capturing reform of limited liability, central bank and incorporation law. Other variables are time invariant. Urbanization controls include the share of the population resident in urban areas and a dummy for suburban counties. See tables B1 and B.2 in the online appendix for further definitions. All specifications are Poisson. Standard errors clustered at county level (236 counties). Fixed effects at the level of 14 districts. All regressions exclude Berlin. Standard errors in parentheses: + $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 7
Economic significance: Factor change in growth rate of credit cooperatives after one standard deviation change in independent variables (last column).

Variable	Coefficient (β_x)	Incidence Rate Ratio (IRR)	Standard deviation of independent variable σ_x	$\beta_{standard}$
Δ grain prices	−38.399	2.107E-17	0.011	**0.66**
Population	0.007	1.007	32.74	**1.25**
farm size	0.141	1.151	4.81	**1.97**
farm size squared	−0.004	0.996	133.61	**0.59**
land inequality	−0.849	0.428	0.172	**0.86**
urbanization	−1.351	0.259	0.211	**0.75**
limited liability	1.592	4.914	(dummy)	(dummy)
central bank	1.389	4.010	(dummy)	(dummy)

Coefficient (β_x): as displayed in the benchmark regression (3), Table 1. IRR: Factor change in co-op growth after 1 unit change in independent variable ($\exp(\beta_x)$). $\beta_{standard}$: Factor change in co-op growth after 1 std. dev. change in independent variable ($\exp(\beta_x{}^*\sigma_x)$).

However, the effect of fractionalization could be obscured, if ethno-linguistic or confessional groups differ by characteristics that are relevant to founding or operating MFIs. For example, an influential literature accords higher human capital formation (Becker and Wößmann, 2009) or entrepreneurial productivity (Hornung, 2014) to Protestants in Prussia. Moreover, ethnic Poles were likely to have received less education than Germans, as ethnically Polish areas received less funding from the Prussian government (Cinirella and Schueler, 2016). In this case, our fractionalization indices could be capturing the effect of moving along the population's skill distribution. In column (3) and (4) we therefore control for the share of Polish speakers and Protestants respectively. In the first case, the positive coefficient on linguistic fractionalization becomes marginally significant. The sign on the share of Polish speakers shows that Polish areas were indeed somewhat less likely to see credit cooperatives develop.



(a) Data outcome and model fit with price volatility

(b) Data outcome and model fit with interest rates

Fig. 3. Model fit over time.

It could also be that ethnic heterogeneity only matters if it is correlated with economic inequality (Alesina et al., 2016). For example, O'Rourke (2007, p. 1375) shows for Irish cooperatives that it is "this coincidence between religious, national and class divisions within Ireland that was crucial for the slow spread of cooperation". On the other hand, in the context of the model in section 2, if the wealthy resort to bank loans, and wealth and ethnicity overlap, the remaining cooperative group could be more homogenous. This could actually lower their internal transaction costs. We tackle these possibilities by interacting linguistic fractionalization and land inequality in column (5) as well as calculating our own measure of religious inequality in column (6).[51] However, neither measure yields significant results. Importantly, none of the other coefficients in the model are affected, including our baseline measure of land inequality. We therefore conclude that there is no consistent role for population heterogeneity in our results.

4.5. Cooperative characteristics and further robustness

Until this point, we have pooled all credit cooperatives in our data set. However, our data does allow us to differentiate cooperatives according to their characteristics, in particular the liability structure and deposit requirements adopted at the end of the sample. We first exclude all cooperatives in counties with an urban population above the median in column (1) of table F5. We do this as urban cooperatives often followed the organizational model of Schulze-Delitzsch rather than the Raiffeisen model, which dominated in rural areas. However, we do not find that this changes the results. Similarly, we do not find in column (2) that only including cooperatives that had opted for unlimited joint liability by 1913 affects our conclusions.[52] The subset of unlimited liability cooperatives can be further decomposed according to the minimum capital contribution required by the cooperative. Reassuringly, we find cooperatives requiring "large" contributions (above 100 Mark) in column (4) to be more affected by the distribution of assets in their county than those at or below this threshold in column (3).[53] The role of price shocks is apparent for both categories.

One remaining concern could be that our results are driven by a few instances of particularly drastic price changes. One episode that stands out in Fig. 1c is the rapid fall in grain prices in the early 1890s, which preceded a large growth episode for credit cooperatives. In column (2) of table F6 we check whether excluding this spike affects our conclusions. Although the coefficient on grain prices turns out to be smaller, we conclude that this episode alone does not drive our story. A related worry could be that this large price shock acted as a structural break, fundamentally altering the relationship between our variables. Yet inspecting columns (3) and (4) does not suggest that significant differences exist before and after the shock, although the effect of land inequality is weaker after the 1890s as noted in section 4.2.

We also check whether our benchmark results are robust to changes in the cross-sectional sample composition. Table F7 follows a Jackknife approach and drops all observations from one of our six provinces in turn. The stability of the coefficients show that our results are unlikely to be driven by outliers.

Finally, we investigate the sensitivity of our results to changes in the econometric specification. Three issues in particular are important. Firstly, one might worry about the large proportion of zero-outcomes in our data. Our Poisson model pooled the zero and non-zero outcomes, while the data generating process for both could be different. One common remedy are zero-inflated models, which allow both processes to differ (Cameron and Trivedi, 2013). We implement these models in column (1) and (2) of table F.8, but do not find that it affects our conclusions. Secondly, we relax our assumption that co-op growth follows a Poisson distribution. Columns (3) and (4) implement a regular and zero-inflated Negative Binomial model respectively. Once again, this does not change our results. Thirdly, we regress our count variable on the natural logarithms of all independent variables in column (5) using Ordinary Least Squares. In column (6), we do the same for the logarithmic transformation of the co-op growth rate. Although these are not our specifications of choice, as they ignore the count nature of the data, it is reassuring that our conclusions do not depend on specific functional forms.

4.6. Model fit, economic significance and counterfactual analysis

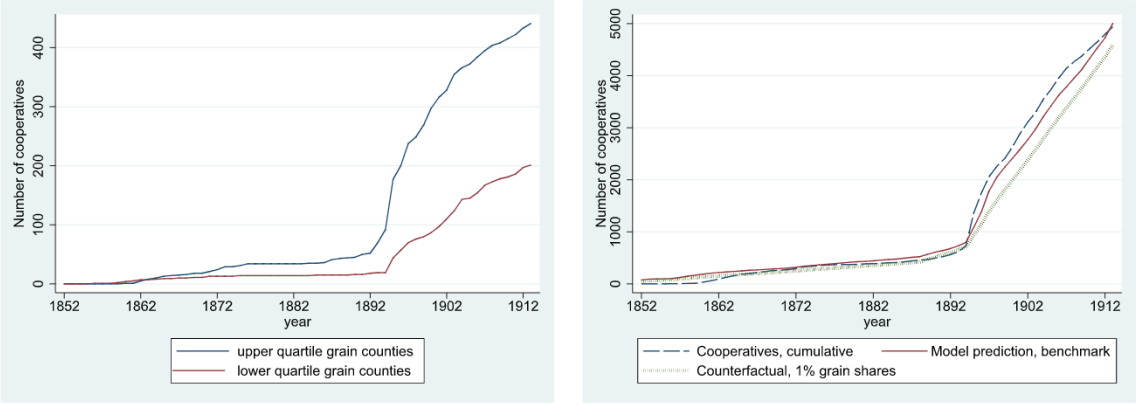
The previous sections have been mainly concerned with investigating the extent to which the data supported the predictions of our theoretical framework. Now we turn to economic significance: to what degree have we actually explained the rise of credit cooperatives?

Our first concern is model fit over time. Fig. 3a plots the data on aggregate growth of cooperatives (dashed line) against the aggregate annual predicted values from the benchmark specification (red solid line). This solid line only reflects price changes and policy dummies. The prediction tracks the data quite closely. Clearly, the policy dummies account for the level differences after the 1890s, highlighting the

[51] We measure the extent to which members of either the Protestant or Catholic confession are overrepresented among land-holding farmers in a county. Data is from the 1882 Prussian occupational census. For groups A, B in county i: $religious\ inequality_i = \left| \frac{Farmers_{A,i}}{Total\ employed\ in\ agriculture_{A,i}} - \frac{Farmers_{B,i}}{Total\ employed\ in\ agriculture_{B,i}} \right|$

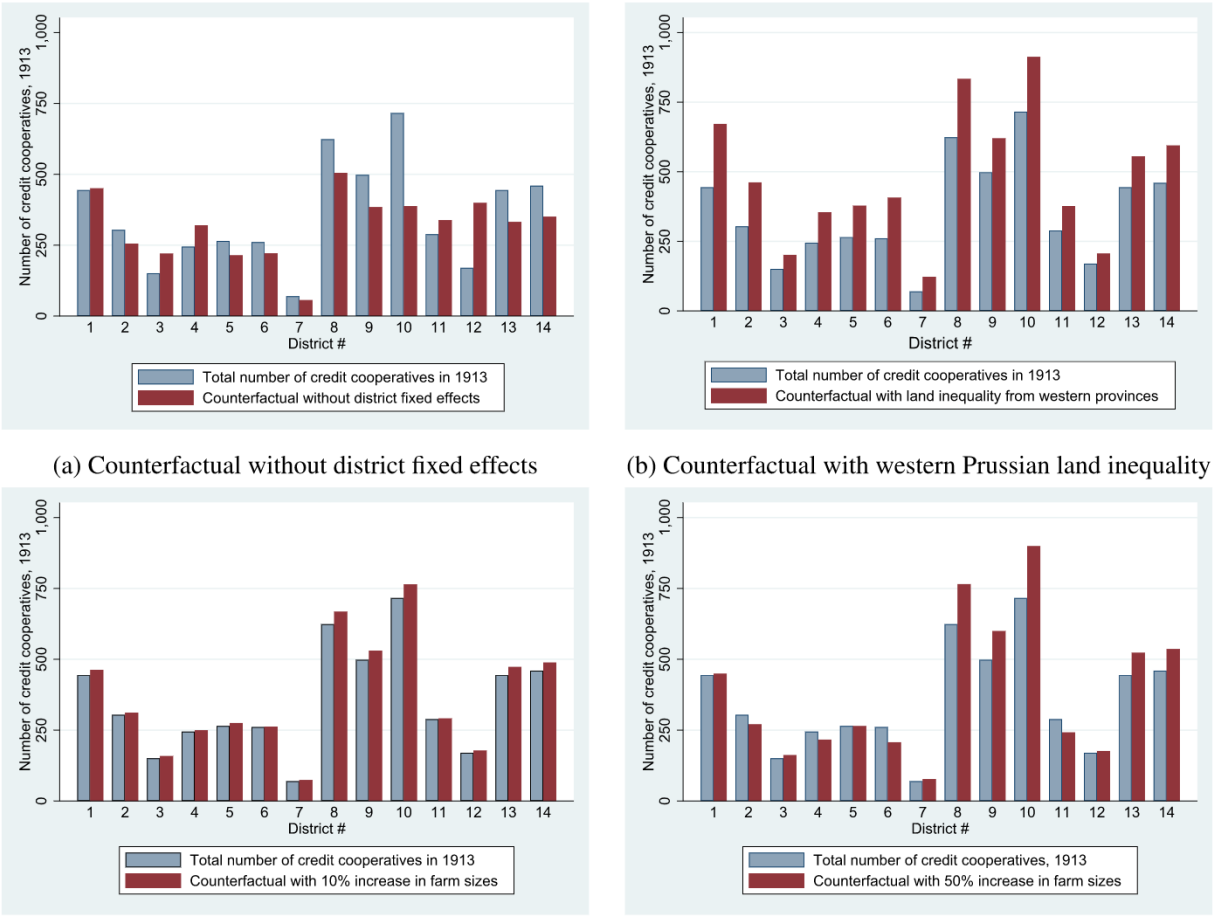
[52] The principle of unlimited liability was largely followed by Raiffeisen cooperatives, and was one of the main distinguishing features of that organisational model.

[53] 100 Mark was equal to roughly 50 days of work at 1913 average wages.



(a) Lowest and highest quartile grain countries (b) Counterfactual grain shares at lowest 1%

Fig. 4. Role of rye and wheat cultivation for cumulative growth of cooperatives: Descriptive statistics and counterfactual.



(a) Counterfactual without district fixed effects (b) Counterfactual with western Prussian land inequality

(c) Counterfactual with 10% farm size increase (d) Counterfactual 50% farm size increase

Fig. 5. Data outcomes and counterfactual scenarios: final number of cooperatives in 1913 at the district level.

importance of the institutional environment for small scale credit. Still, the grain price variable is able to replicate many turning points in the data. Would including more time-varying determinants further improve model fit? The third line in panel 3a plots the benchmark specification after the inclusion of price volatility. However, price volatility does not add much explanatory power. In panel 3b we substitute price volatility for interest rates, which creates a very good model fit. We conclude that we are able to largely explain the development of rural microcredit with reference to only prices and interest rates.

Our price variable consists of the interaction between temporal price changes and local land use structure. How important is local variation in land use in determining the final number of cooperatives? In Fig. 4a we plot the rapid increase in the cumulative number of cooperatives for the counties in the highest quartile of the distribution for both rye and wheat cultivation. This is in contrast to the sluggish evolution of co-ops in low-grain countries. This descriptive evidence suggests that differ-

ences in cross-sectional reliance on grain may be an important determinant of the final number of cooperatives. Fig. 4b engages in a counterfactual experiment by setting local rye and wheat shares to the first percentile in their respective distributions. The counterfactual evolution of cooperatives is slower than the benchmark model prediction, although cross-sectionally induced differences seem relatively small compared to high average growth rates over time.

Having established the explanatory power of our model over time, our next interest is the relative importance of individual variables. In Table 7 we list all statistically significant coefficients from our benchmark specification. The coefficients are converted into a count data equivalent of standardized β-coefficients (last column).[54] Grain price movements clearly had a large impact: One standard deviation increase in relative grain prices decreases the growth rates of credit cooperatives by a factor of 0.66. Asset size plays a large role too-a one standard deviation sized increase in farm sizes almost doubles co-op growth. The role of population is more muted, and the increase in the co-op growth rate to an increase in population is underproportional. Similarly, asset inequality seems to have exercised a modest influence. However, this could partially be attributed to the relatively limited variation in inequality in our sample. To examine this, we turn to counterfactual simulations.

Graph 5a plots the final number of credit cooperatives in 1913 for all of our 14 districts.[55] By virtue of the district fixed effects, this is equal to the predicted values of the benchmark regression by district (not shown). The first counterfactual consists of setting the fixed effects to 0, and shows that a large share of the variation is explained by the remaining covariates.

Given the relatively limited variation in inequality in the sample, we need to consider a "radical" policy experiment to see the effect of changes in this variable. In graph 5b, we set land inequality to the level prevailing in the western provinces of Prussia, which is on average 2 standard deviations lower than our sample mean. This leads to a non-trivial counterfactual change in the number of credit cooperatives. In particular the most unequal districts (comprising the province of Posen, see table D1) would have seen a 50% increase in the number of micro credit institutions.

We have shown asset size to be both statistically and economically highly significant. However, the direction of the economic effects is hump-shaped. We illustrate this using a "moderate" and a "radical" policy experiment. Both experiments rely on the distribution of assets within the county (inequality) remaining unchanged, while the mean is allowed to change. In the moderate experiment (graph 5c) we increase average farm size by 10% in all counties. The result is a modest counterfactual increase in the number of cooperatives in all districts, suggesting that most counties were below the hump. In the radical experiment (graph 5d), we increase farm sizes by 50%. In this case, the negative impact of the quadratic term on farm size dominates in some regions. While districts marked by small farms gain a lot, others see a counterfactual decrease in cooperative activity. As in theory, large increases in assets causes potential borrowers to substitute bank for small-scale cooperative credit.

5. Conclusion

Starting from a simple theoretical framework, we have derived a number of factors that explain the development of rural micro credit both across time and space. We provided empirical evidence on the development of micro finance in a free market benchmark case without public ownership, subsidization and under a minimal regulatory environment: the six eastern provinces of Prussia between 1852 and 1913. Our results show that asset sizes, asset inequality, interest rates, and relative prices of agricultural staples explain a large share of the variation in credit cooperatives. We argued that these price movements can be seen as one driver of structural transformation, and that micro credit can potentially play a part in rural adaption to this transformation.

We briefly point to two extensions and one implication of our findings.

Firstly, our approach has explained the development of micro credit institutions as partly depending on complementary financial infrastructure from banks. However, we have not accounted for informal village money lenders, partly because such data is lacking for our period. A recent literature studies the impact of competition between MFIs and money lenders on market outcomes, including interest rates (Mookherjee and Motta, 2016; Demont, 2016). It would be interesting to extend our results in this dimension and examine how the entry decisions of MFIs depends on existing informal money lending.

Secondly, although we have considered economic inequality and ethnic heterogeneity, we have abstracted from many other individual characteristics of potential borrowers. As Ghatak (1999) has classically shown, information about individual-level riskiness can influence group formation through assortative matching. Future research could investigate to what extent regional differences in information dissemination, network structures and group cohesion impact the dispersion of micro credit institutions.

Lastly, we discuss the implications of our focus on long-run structural transformation for understanding the impact of micro credit. We have presented a scenario with two production sectors available to farmers, with sector profitability trending in opposite directions. It follows, firstly, that micro credit uptake will only occur in a limited time frame, namely when sector profits are close enough for switching to be feasible. Secondly, in such a setting there is likely to be a market demand for medium term loans rather than short term financing. Thirdly, if microcredit allows farmers to exit a declining sector, a comparison between their pre- and post credit incomes will not necessarily reveal an "impact" of microfinance on incomes. A more relevant outcome metric would be changes in the sector of economic activity. In this sense, microfinance may be better understood as aiding adaption to a changing environment (Goodspeed, 2016) or as broadening choices (Banerjee et al., 2015), rather than as a tool allowing rural populations to escape poverty.

CRediT authorship contribution statement

Marvin Suesse: Conceptualization, Methodology, Validation, Writing - original draft, Formal analysis, Writing - review & editing. **Nikolaus Wolf:** Conceptualization, Methodology, Resources, Validation, Writing - original draft, Writing - review & editing.

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Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.jdeveco.2019.102429.

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[54] $\beta_{standard}$ shows the factor change in the co-op growth rate as a response to a change in the independent variable. A value close to 1 thus signifies a small effect.

[55] District numbers correspond to those defined in table D1.

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