Credit Expansion and Housing Cycle

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Abstract

During the 1999-2009 U.S. housing cycle, two opposing empirical facts present a puzzle: the correlation between income growth and mortgage growth is negative across ZIP codes within metropolitan areas (some argue for the 'credit expansion' view) but positive across metropolitan areas (others argue for the 'speculation' view). First, I show that the cross-metropolitan phenomenon, in fact, is also consistent with the 'credit expansion' view: by an instrumental variable approach, I show that net export growth across metropolitan areas causes both income growth and credit expansion in mortgage growth, which eventually leads to the housing cycle. I also design five tests with detailed measures to show that credit expansion rather than speculation plays the dominant role. Second, I develop a new theoretical model that, for the first time in the literature, reconciles the above two seemingly opposing empirical facts in the 'credit expansion' view. Third, my theoretical model generates new predictions of 'double differences' in the literature: the differential stronger boom and bust cycle in mortgages (and house prices) in low-income ZIP codes than in high-income ZIP codes within metropolitan areas is more pronounced in high net-export-growth metropolitan areas. I provide empirical causal evidence for these new predictions, further supporting the 'credit expansion' view.

Keywords: housing cycle, credit supply, private-label mortgages
JEL Classification: G01, G21, R31

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"Ten years after the financial crisis, the central question of what explains the rise and fall in house prices remains unresolved."

– Griffin, Kruger, and Maturana, *JFE, 2021*

1 Introduction

The U.S. housing price boom and bust in the 2000s are unprecedented. In the bust period, U.S. households experienced the deepest recession since the Great Depression with widespread mortgage defaults (Keys, Mukherjee, Seru, and Vig, 2010; Mayer, Pence, and Sherlund, 2009), large consumption drop (Kaplan, Mitman, and Violante, 2020; Mian, Rao, and Sufi, 2013), and huge unemployment rise (Hoynes, Miller, and Schaller, 2012; Mian and Sufi, 2014). Understanding the cause of such a huge housing boom and bust is crucial for understanding the economic connections among credit, the housing cycle, and economic activities. It is also helpful for the design of housing policy and the regulation of mortgage credit.

In the literature, there are two competing views regarding the cause of the above housing cycle: supply-side excess credit expansion by lenders and demand-side speculation by borrowers. These two competing views are based on two seemingly opposing empirical facts: the correlation between income growth and mortgage growth is negative across ZIP codes within metropolitan areas but positive across metropolitan areas. Leading by Mian and Sufi (2009), one group of scholars argues that the negative correlation supports the "credit expansion" view. However, started by Adelino, Schoar, and Severino (2016), another group of scholars argues the positive correlation supports the "speculation" view.

To preview, this chapter reconciles two empirical facts within the "credit expansion" view by empirical tests, a theoretical model, and empirical tests supporting a new prediction by the model. My core finding is that credit expansion rather than speculation plays the dominant role in causing the 1999-2009 U.S. housing cycle. Therefore, the policy implies that the new regulatory design to prevent the same housing cycle shall focus more on the lenders. In the following part of the introduction, I will describe the research design, major findings, and contribution.

Research Design Researchers face several difficulties in identifying the causal evidence of credit expansion on housing prices in the United States. First, any identification strategy needs cross-section differences in both credit expansion and housing prices, which in turn requires the treatment variable (incentive)[1] to have enough area coverage (U.S. metro), time coverage (99-09), and geographic variation. Second, the geographic variation in the treatment variable must persist over

[1]To be precise, the 'treatment variable' here refers to the instrumental variable in an IV strategy, the running variable in a regression discontinuity design, and the treatment variable in a difference-in-difference setting. For example, Di Maggio and Kermani (2017) uses the interaction between 2004 OCC preemption and the state-level anti-predatory lending law as the treatment variable in their difference-in-difference study.

the housing boom period since the mortgage is a long-term financial decision. In other words, the treatment variable must provide persistent incentives to induce stronger credit expansion in certain areas than others throughout the entire boom period. Short-run shocks, such as weather (e.g., rain, extreme temperature, and wildfires), are inadequate. Third, beyond the variable nature, shocks to the treatment variable are required to constitute a reasonable identification strategy. Fourth, the underlying economic theory should explain why the treatment variable can induce stronger credit expansions in the boom period but not in the prior period. Such a theory requires a deep understanding of the incentives of different types of participants, the mortgage market structure, and their time-varying nature.

The research design in my study addresses the above challenges in three parts. In the first part, I operationalize the central idea of the "economic base theory" (Tiebout, 1962) and construct a treatment variable that captures the long-term incentive of credit expansion: metropolitan exposure to net export growth of manufacturing industries (hereafter "net export growth"). Regional economics defines the "economic base sector" (tradable sector) as the economic activities that a local area provides for the world outside, thus bringing wealth to the local area. Most money will be reused locally among the "non-base sector" (nontradable sector) via a multiplier effect. According to this theory, the tradable sector growth is the key driving force for local economic growth and house prices in the long term. Therefore, credit expansion would be stronger in areas with stronger tradable sector growth due to the (1) higher foreclosure price of the house and (2) higher income growth in the future that can be recoursed by lenders given default. Ideal measurement requires census-style data covering the accounting data of all firms in the tradable sector, which is unavailable. Instead, I proxy the composite (share) with manufacturing employment data at the industry-by-metropolitan level. In addition, I use the large time-series change in net export growth in manufacturing in the U.S. as a proxy for the relative growth (shift) at the industry level. Aggregating these two proxies (shift and share) can give us a good measure of the relative growth of the local tradable sector across metropolitan areas. By this construction, net export growth has enough area coverage (U.S. mainland) and enough time coverage (91-09). In addition, its geographic variation arises from the fact that the local manufacturing sector tends to cluster within several related industries due to internal and external economies of scale. Further, the persistence feature of net export growth arises from two dimensions. At the industry level, both comparative advantages (due to differences in technology and endowment) and horizontal specialization (due to economies of scale) across countries make trade patterns persistent over time. At the local level, the formation of industry clusters (involving establishment building and work migration) with huge costs is unlikely to change rapidly in a short time. At the individual level, job reallocation across different industries is very costly.

In the second part of my research design, I use the instrumental variable approach developed by

Feenstra, Ma, and Xu (2019) as my identification strategy. Feenstra, Ma, and Xu (2019) develop their IVs from a general equilibrium setting that can isolate the exogenous part of U.S. imports and exports. Intuitively, their IVs capture the exogenous parts of net export growth due to (1) rising world demand reflected in U.S. export growth, (2) rising world supply reflected in U.S. import growth, and (3) tariff changes. They also use high-dimensional fixed effects to remove the potentially endogenous parts and pre-determined parts: (1) U.S. industry-by-year supply shocks in exports, (2) U.S. industry-by-year demand shocks in imports, and (3) pre-determined bilateral distance between U.S. and partner countries. They construct IVs for exports and imports separately, and I combine both as an IV for net export growth.

In the third part of my research design, I build an economic model to show that credit expansion alone can explain the two seemingly opposing empirical facts: the correlation between income growth and mortgage growth is negative across ZIP codes within metropolitan areas but positive across metropolitan areas. This model is also consistent with the viewpoint that net export growth cannot induce house price growth in the prior period (91-99). A key legal constraint is that government-sponsored enterprise mortgages cannot consider differences in regional economic conditions (growth) (Hurst, Keys, Seru, and Vavra, 2016). But private-label mortgages can. Since "global saving glut" (Bernanke, 2005; Bernanke et al., 2007), securitization innovation (notably the Copula approach by Li (2000) (Salmon, 2012)), mortgage market deregulation (Di Maggio and Kermani, 2017; Lewis, 2023), and political push (Mian, Rao, and Sufi, 2013) all happened after 1999, private-label mortgages had high mortgage rates between 1991 and 1999 and only occupied a modest market share. In contrast, government-sponsored enterprise mortgages dominated the mortgage market with low rates due to the government's implicit guarantee and economies of scale. Consequently, without aggregate credit expansion, even high net-export-growth areas cannot experience a boom and bust cycle.

Findings I illustrate my findings in five parts. My first three findings together show that the "speculation" interpretation regarding the positive correlation between income growth and mortgage growth across metropolitan areas has an omitted variable problem. The "speculation" view omits regional economic growth, for which I proxy it with metropolitan exposure to net export growth of manufacturing industries. The net export growth causes income growth and credit expansion in mortgages across metropolitan areas. In my fourth part, I use a model to show that credit expansion can reconcile the above two empirical facts, thus addressing the potential concern regarding the internal conflicts between the two empirical facts.

In the first part, I document a new empirical fact: a stronger housing cycle in the high net-export-growth metropolitan areas than the low net-export-growth ones in the U.S. between 1999 and 2009, as shown in the following Figure (1).[2] From 1999 to 2005, the house price growth in the

[2] "High net-export-growth metropolitan areas" means those with high exposure to manufacturing industries' net

highest quintile of net-export-growth areas was 73% but only 41% in the lowest quintile.[3] From 2007 to 2009, the house price drop in the highest quintile is 22% but only 13% in the lowest quintile. In sharp contrast, there is little difference in house price growth between the highest and lowest quintile areas in the prior period (91-99).

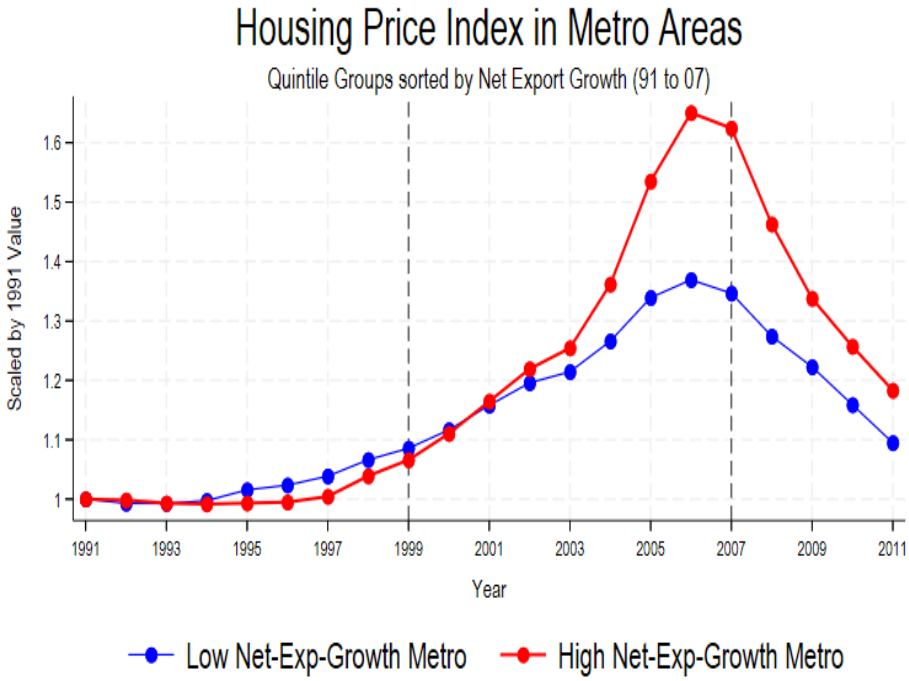


Figure 1: Housing Price Index (91-11) across Metropolitan Areas

In the second part, using the instrumental variable approach by Feenstra, Ma, and Xu (2019), I provide causal evidence that, induced by net export growth, credit expansion in mortgages causes the housing price boom and explains the bust across metropolitan areas.[4] Specifically, growth in the private-label mortgage (non-jumbo) in the boom period (99-05) is much stronger in the high-net-export-growth metropolitan areas, causing a much stronger house price boom (99-05) and bust (07-09) than those in the low net-export-growth metropolitan areas.

Please note that "causal evidence" in this paper only means that I find an incentive (net export growth) that induces credit expansion in private-label mortgages to be much stronger in the high

export growth. For simplicity, I use this terminology throughout this paper.

[3]This differential increase rate of 32% is very large compared to the commonly documented within-metropolitan across-zip code differences in house price. For example, in Figure 1 in Griffin, Kruger, and Maturana (2021), for the within-metropolitan across-zip code house price growth based on four credit supply variables and three speculation variables, the differential increase rates between the top and bottom quartiles are all below 20%, between the end of 2002 to the end of 2006.

[4]Since the literature has already documented that net export growth increases household income (Barrot, Loualiche, Plosser, and Sauvagnat, 2022; Minot and Goletti, 1998), I leave the verification test of net export growth on income growth to the appendix Section A.5.

net-export-growth areas in the cross-section. I do not find the incentives that caused the aggregate credit expansion in the early 2000s. For the causes of the aggregate credit expansion in the boom period (1999-2005), the literature has documented evidence from international capital flow ("global saving glut" mainly 2003-2007 by Bernanke (2005); Bernanke et al. (2007)), financial innovation in securitization (notably the Copula formula by Li (2000) (Salmon, 2012)), mortgage market deregulation (2004 preemption of national banks from the state-level anti-predatory lending law by the Office of the Comptroller of the Currency (Di Maggio and Kermani, 2017), 2005 Bankruptcy Abuse Prevention and Consumer Protection Act (Lewis, 2023)), and political campaign (2002-2007 mortgage industry campaign contributions (Mian, Sufi, and Trebbi, 2013)). Therefore, my choice of 1999-2005 as the boom period tries to include all major events above and matches the mortgage boom period commonly used in the literature (see Griffin, Kruger, and Maturana (2021) for a review).

In the third part, I use five tests to show that credit expansion rather than speculation plays the dominant role in causing the house price cycle above. In my first test, I show that net export growth causes private-label mortgage (non-jumbo) growth but not government-sponsored enterprise mortgage growth during the boom period (1999-2005). This test helps address the speculation (demand) view because demand does not show up in the chapter mortgage type (GSEM). My second test helps to validate the exclusion restriction of my IV strategy: net export growth causes house price growth in the boom period (99-05) only because it causes growth in private-label mortgages (non-jumbo). Specifically, in a prior period (1991-1999) without substantial credit expansion, net export growth does not cause house price growth because it does not cause growth in private-label mortgages (non-jumbo). In other words, this placebo test shows that net export growth cannot directly cause house price growth via the demand channel without credit expansion in mortgages. In the following three tests, I use a specific measure of pure speculation: "non-owner-occupied" private-label mortgages for home purchases (Gao, Sockin, and Xiong, 2020). Inspired by this measure, I use "owner-occupied" private-label mortgages for home purchase as a measure of pure credit expansion. First, I show that pure credit expansion induced by net export growth can explain more than half of the pure cross-metropolitan variation in speculation. The residuals give us a measure of credit-independent speculation. My second test shows that, compared to credit-independent speculation, pure credit expansion is dominant in explaining the house price boom. Specifically, pure credit expansion explains twelve times the house price growth that speculation explains. My last test focuses on the prior period (91-99) with substantial net export growth but no credit expansion. I show that, without credit expansion, pure speculation does not increase as a response to the net export growth. Taken together, the results of my five tests are more consistent with the notion that credit expansion rather than speculation plays the dominant role in causing the house price boom and bust. My argument against demand does not mean demand does not play any role. Instead, I only mean that credit expansion plays the dominant role compared to

speculation.

In the fourth part, I build a model to illustrate that credit expansion alone can reconcile the two seemingly opposing empirical facts: the correlation between income growth and mortgage growth is negative across ZIP codes within metropolitan areas but positive across metropolitan areas. The basic intuition of the cross-metropolitan fact can be illustrated in the following Figure (2). Conceptually, U.S. metropolitan areas can be sorted into high and low net-export growth areas. Higher net export growth results in higher employment growth, higher household income growth, and higher population growth in the high net-export-growth area than in the low net-export-growth area. These differences give the mortgage borrowers in the high net-export-growth area two advantages: (1) higher foreclosure price of house given default [5] and (2) higher income growth that can be recoursed by lenders given default.

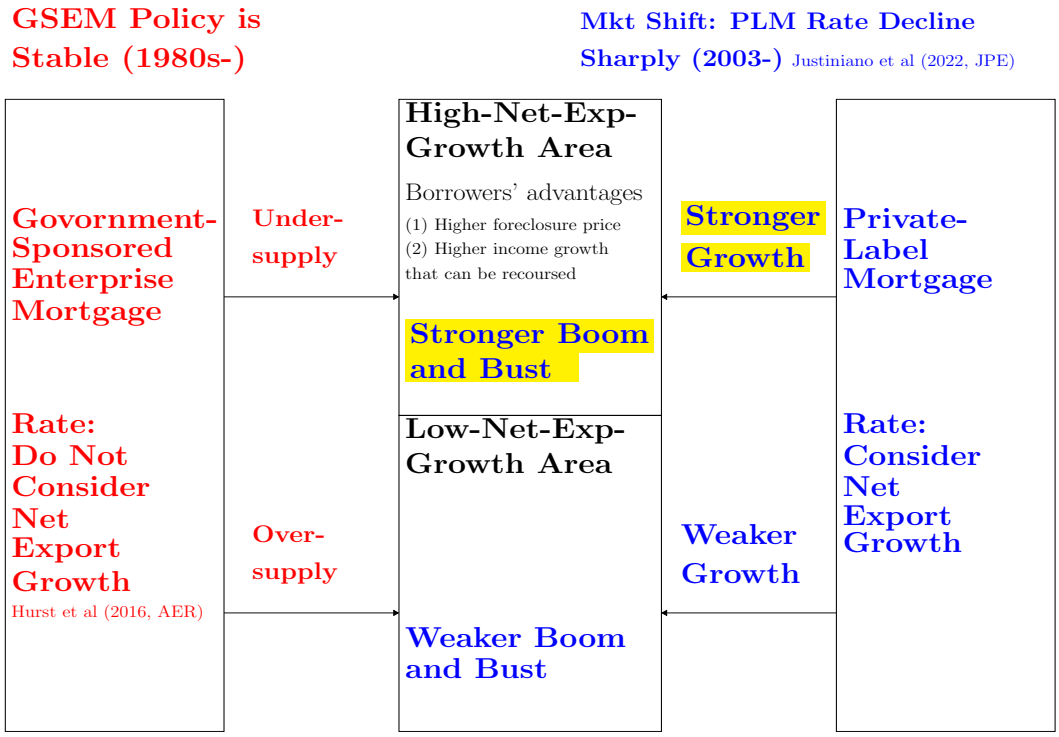


Figure 2: Model Intuition

A key legal constraint is that government-sponsored enterprise mortgages (GSEMs) cannot consider differences in regional economic conditions (growth) in setting up mortgage rates (Hurst, Keys, Seru, and Vavra, 2016). They show that GSEMs essentially tax areas with lower default rates while subsidizing areas with higher default rates. In the spirit of their argument, I argue that,

[5] In the empirical literature, there is evidence that employment growth, household income growth, and population growth (including migration) can push up housing demand and then housing price, especially in the long term (Olsen, 1987).

due to the legal constraint, government-sponsored enterprise mortgages essentially relatively over-supply mortgage credit to the low net-export-growth metropolitan areas and relatively under-supply mortgage credit to the high net-export-growth metropolitan areas. In sharp contrast, private-label mortgages consider these differences when setting mortgage rates. For a long time before the boom period (1999-2005), due to the government's implicit guarantee and economies of scale, government-sponsored enterprise mortgages dominated the mortgage market because of their low rates. Consequently, private-label mortgages were not able to correct such misallocation.[6] However, documented by Justiniano, Primiceri, and Tambalotti (2022), there is a sharp and persistent credit expansion in the private-label mortgages that started in 2003 summer.[7] Using loan-level data and a regression model, they identify a sharp and persistent decrease in the spread between private-label mortgages and 10-year treasury yield in 2003 summer. I model this empirical fact as a declining funding cost in private-label mortgages. My model predicts that private-label mortgages show a stronger increase in high net-export-growth metropolitan areas because of the above two advantages in mortgage borrowing (higher foreclosure price and higher borrower income growth) compared to low net-export-growth metropolitan areas. This differential increase in private-label mortgages in high net-export-growth areas results in a stronger housing boom and bust cycle.

Next, the basic intuition of the within-metropolitan fact is as follows. When the funding cost of private-label mortgages declines, lenders can extend cheap credit to low-income borrowers traditionally excluded from the mortgage markets due to usury law and credit rationing. Since households with sharp rising wealth tend to self-relocate to richer ZIP codes and households with sharp declining wealth tend to self-relocate to poorer ZIP codes, income growth in poorer ZIP codes is usually lower than in richer ZIP codes, even if income growth distribution is the same across all occupations, industries, and initial wealth level. Thus, the negative correlation between income growth and mortgage growth within metropolitan areas means that poorer ZIP codes see higher mortgage growth.

Then, my model reconciles the above two empirical facts. Specifically, due to the differences in net export growth, poorer ZIP codes in the high net-export-growth areas experience relatively higher private-label mortgage growth than those in the low net-export-growth metropolitan areas. Since most credit expansion in private-label mortgages happened in the poorer ZIP codes in each metropolitan area, the correlation between income growth (aligned with net export growth) and mortgage growth is positive across metropolitan areas.

In the fifth part, my model generates two additional predictions of "double differences" in mortgage and house price growth. I create a low-minus-high factor: I use the mortgage (and house

[6] Estimates show that the spread between government-sponsored enterprise mortgages and otherwise similar jumbo loans (purchased by private issuers) are, on average, between 15-40 basis points between 1996 and 2006 (see Sherlund (2008) and its summary of the literature).

[7] Justiniano, Primiceri, and Tambalotti (2022) argues that this rate drop likely reflects mispricing, as shown in the subsequent increasing default rate.

price) growth in the low-income ZIP codes minus the one in high-income ZIP codes within the same metropolitan area. Consistent with the within-metro fact (Mian and Sufi, 2009), this low-minus-high factor in mortgages (and house prices) is positive in the boom period and negative in the bust period. The first prediction is that net export growth in the boom period causes the boom and bust in the low-minus-high factor in private-label mortgage (non-jumbo) growth. And the second prediction is that the low-minus-high factor in private-label mortgage (non-jumbo) growth in the boom period causes the boom and bust in the low-minus-high factor in house price growth.

To summarize the model, I show that credit expansion in private-label mortgages induced by net export growth alone can reconcile empirical facts both within and across metropolitan areas. In addition, causal evidence confirms the model-based predictions of "double difference", thus reinforcing the "credit expansion" view.

I conduct multiple robustness tests for my main conclusion that private-label mortgage (non-jumbo) growth in the boom period causes house price boom and bust. First, I consider the state-level anti-predatory lending law and the 2004 preemption of national banks by the Office of the Comptroller of the Currency (Di Maggio and Kermani, 2017). Although preempted states experience a stronger housing boom and bust cycle caused by credit expansion, my main conclusion still holds in all states. Second, I test state-level differences in recourse law (Ghent and Kudlyak, 2011). Even though non-resource states experience a stronger housing bust caused by credit expansion, my main conclusion is robust for all states. Third, I investigate state-level differences in the judicial requirement of foreclosure laws (Mian, Sufi, and Trebbi, 2015). Though non-judicial states experience a stronger house price bust caused by credit expansion, my main conclusion still holds in all states. Fourth, my main conclusion still holds after controlling the sand-state dummy. Fifth, my main conclusion is robust to state capital gain tax.

Contribution to the Literature I illustrate my contribution to the literature in five parts. First, I document a new empirical fact in the literature: a stronger housing boom and bust cycle in the high net-export-growth metropolitan areas relative to the low net-export-growth ones in the U.S. between 1999 and 2009. Prior literature mainly focuses on the cross-state legal differences in mortgage markets (Choi, Hong, Kubik, and Thompson, 2016; Di Maggio and Kermani, 2017; Gao, Sockin, and Xiong, 2020; Ghent and Kudlyak, 2011; Mian, Sufi, and Trebbi, 2015).

Second, my paper uses empirical tests and a theoretical model to show that credit expansion alone can reconcile the two seemingly opposing empirical facts: the correlation between income growth and mortgage growth is negative across ZIP codes within metropolitan areas but positive across metropolitan areas. Leading by Adelino, Schoar, and Severino (2016), a group of studies argue that the positive correlation across metropolitan areas is more consistent with the "speculation" view. However, I show that their omitted variable, net export growth, drives both income growth and credit expansion in private-label mortgages. Such credit expansion eventually results

in a differentially stronger boom and bust cycle in the high net-export-growth areas in the U.S. between 1999 and 2009. In addition, I provide a simple theoretical model to show that credit expansion alone can reconcile the two seemingly opposing empirical facts. The model-based predictions of "double differences" get confirmation by causal evidence, reinforcing the "credit expansion" view by Mian and Sufi (2009).

Third, my paper makes unique contributions to the literature in five dimensions on the causal impact of credit expansion on the 1999-2009 U.S. housing cycle. First, my paper explains the differential housing cycles across metropolitan areas (M.A.), a new dimension in the literature. Second, my empirical design builds on the "economic base theory" that the tradable sector determines the local housing market in the long term. Thus, net export growth determines the long-term direction of credit expansion in private-label mortgages. Third, using an instrumental variable approach from international economics, I provide the causal evidence that credit expansion in the boom period (1999-2005) causes the house price boom (1999-2005) and explains the bust (2007-2009) in the cross-section. A critical advantage of the IV approach by Feenstra, Ma, and Xu (2019) is that this method can cover the entire period of credit boom consisting of all major events of securitization innovation, international capital flow, financial deregulation, and political campaign. In comparison, some other causal studies mainly focus on one single event (Di Maggio and Kermani, 2017; Lewis, 2023). As for the causes of the aggregate credit expansion, the literature has documented evidence from securitization technology innovation (notably the Copula formula by Li (2000) (Salmon, 2012)), international capital inflow ("global saving glut" mainly 2003-2007 by Bernanke (2005); Bernanke et al. (2007)), financial deregulation (preemption of national banks from the the anti-predatory lending law at state level by the Office of the Comptroller of the Currency in 2004 (Di Maggio and Kermani, 2017), 2005 Bankruptcy Abuse Prevention and Consumer Protection Act (Ganduri, 2023; Lewis, 2023)), and political campaign (2002-2007 campaign contributions by the mortgage industry (Mian, Sufi, and Trebbi, 2013)). Fourth, I use a model to distinguish the decision process by government-sponsored enterprise mortgages and private-label mortgages based on a key legal difference: government-sponsored enterprise mortgages cannot consider differences in regional economic conditions. This model predicts that private-label mortgages, rather than government-sponsored enterprise mortgages, drive the differential housing cycles across metropolitan areas. Most other papers do not distinguish these two types of mortgages.[8] My empirical tests verify this prediction. Fifth, causal evidence supports the model-based prediction of "double differences", thus bringing a new empirical fact to the literature.

The above five dimensions distinguish my paper from other papers on the causal evidence of credit expansion. The work most closely related is Di Maggio and Kermani (2017). They

[8]The only two exceptions that distinguish the role of private-label mortgages and government-sponsored enterprise mortgages are Justiniano, Primiceri, and Tambalotti (2022); Mian and Sufi (2022). However, they do not show the irrelevance of government-sponsored enterprise mortgages to the differential housing cycle across metropolitan areas.

exploit the OCC's preemption of national banks (rather than state-chartered depository institutions and independent mortgage companies) from state antipredatory-lending laws (APL) in 2004 and onward. They show that, compared to other states, preempted states experienced stronger growth in credit expansion in mortgages by national banks, house prices, and nontradable employment between 2004 and 2006. These outcomes experience a sharper decline during 2007-2010. My paper differs from theirs in five angles. First, in the cross-section, I explain the cross-metro housing cycles while they explain the cross-state housing cycles. Second, my empirical design builds on the "economic base theory" so that net export growth captures the long-term incentives of credit expansion in mortgages. One advantage of the instrumental variable approach is that it covers periods with all major events related to the housing cycle. In comparison, their paper captures a single legal change as the incentive, thus ignoring securitization innovation (Salmon, 2012), 2005 Bankruptcy Abuse Prevention and Consumer Protection Act (Lewis, 2023), and other changes that happened before 2004. Third, they cannot explain why states without the anti-predatory lending law did not experience a housing cycle in other periods, such as 1991-1999. On the contrary, my model is general enough to be consistent with the viewpoint that no such housing cycle could happen between 1991 and 1999 (please see more details in the model intuition). Fourth, while both papers provide causal inference, my research design can do more. Specifically, my IV strategy can decompose the household speculation (non-owner-occupied private-label purchase mortgages) and show (with three tests) that credit expansion is a necessary condition for speculation. Fifth, I take advantage of a key legal constraint to distinguish the roles of government-sponsored enterprise mortgages and private-label mortgages while they separate national banks from other mortgage institutions by the OCC's preemption.

Two other papers use different approaches and focus on different scopes from mine. Favara and Imbs (2015) exploit the 1994-2005 state-level branching deregulation and show that early-deregulating states experience stronger growth in mortgage and house prices. However, they admit in Figure 4 that deregulation cannot explain the growth in house prices beyond 2002. And they do not study the period from 2005 to 2009, a crucial episode in the housing cycle. Adelino, Schoar, and Severino (2012) use exogenous changes in the conforming loan limit as an instrument for lower financing costs and show that easier access to credit significantly increases house prices.

In the literature on the housing cycle, four interesting narratives argue that credit expansion causes housing boom and bust. First, Mian, Rao, and Sufi (2013); Mian and Sufi (2009) argue that the sharp increase in subprime mortgages leads to the housing cycle. Second, Anenberg, Hizmo, Kung, and Molloy (2019); Mian and Sufi (2022) argue that easier securitization increases the mortgage credit, eventually leading to the housing cycle. Third, Keys, Mukherjee, Seru, and Vig (2010); Keys, Piskorski, Seru, and Vig (2013); Purnanandam (2011) provide evidence that securitization lowers lenders' incentive to screen and subsequently results in poor mortgage performance. Fourth,

Griffin and Maturana (2016a,b); Mian and Sufi (2017); Piskorski, Seru, and Witkin (2015) support that financial fraud induced by moral hazard plays a vital role in mortgage expansion and subsequent crisis. However, most of the above papers use a specification that explores within-metropolitan cross-zip code variation. As pointed out by Griffin, Kruger, and Maturana (2021), this method suffers from endogenous concern. In particular, credit expansion might arise due to housing demand or speculative demand. More importantly, they can not explain the large across-MA variation in the house price cycle, which is almost the same magnitude as within-MA across-zip code variation in the literature (Adelino, Schoar, and Severino, 2016). Unlike their specification, my regression specification and IV strategy explore the across-MA variation in housing prices to obtain causal evidence using an instrumental variable approach from international economics.

Fourth, I design three empirical tests to show that credit expansion is the necessary condition of speculation, measured by non-owner-occupied purchase mortgages. Mian and Sufi (2022) shows some evidence that credit expansion by private-label mortgages can explain various types of speculation. But they do not compare the explanatory power of pure credit expansion vs. credit-independent speculation on housing price boom as I do. More importantly, I further show that, without credit expansion, speculation does not increase in response to economic growth in the prior period (91-99).

Fifth, I contribute to the literature on housing policy regarding government-sponsored enterprise mortgages. Hurst, Keys, Seru, and Vavra (2016) documents that government-sponsored enterprise mortgages do not consider regional risk in mortgage rate setting due to a legal constraint, thus resulting in regional welfare transfer. Based on their paper, I further illustrate that this constraint implies only private-label mortgages can experience stronger growth and potentially overshot in the high net-export-growth areas when the funding cost of private-label mortgages declines sharply, a fact documented by Justiniano, Primiceri, and Tambalotti (2022). To the best of my knowledge, I am the only paper that relates this legal constraint of government-sponsored enterprise mortgages to the cross-metropolitan housing cycle.

2 Research Design

In this section, I illustrate empirical research design. First, I measure the fundamental incentive for mortgage credit by operationalizing the central ideal of the "Economic Base theory" by Tiebout (1962). I use metropolitan exposure to net export growth of manufacturing industries (hereafter "net export growth") as a proxy for the growth of the local tradable sector, which, by theory, is the fundamental force of long-term house price growth. To overcome the endogeneity issue of OLS, I use an instrumental variable approach by Feenstra, Ma, and Xu (2019) as an identification strategy.

2.1 Operationalize the "Economic Base Theory"

In this subsection, I operationalize the central idea of "economic base theory" by using metropolitan exposure to net export growth of manufacturing industries (hereafter "net export growth") as a proxy for the growth of the local tradable sector. This net export growth measure can fulfill the requirement of the treatment variable in the introduction: enough area coverage (U.S. mainland), enough time coverage (91-09), enough geographic variation, and a feature of persistence.

To capture the incentive to credit expansion in mortgages, I revisit the "economic base theory" in regional economics. Regional economics defines the "economic base" (tradable sector) as the economic activities that a local area provides for the world beyond its boundaries (Nijkamp, Rietveld, and Snickars, 1987; Tiebout, 1962). Thus, the tradable sector brings wealth into the local area, and most of the money will be reused locally via a multiplier effect through the nontradable sector. Based on this theory, the tradable sector is the most important driving force for local economic growth, land use, and house prices in the long term (Ling and Archer, 2017; Nijkamp, Rietveld, and Snickars, 1987; Thrall, 2002). Therefore, the tradable sector growth can predict demand-side factors that shape the long-term growth of housing prices, such as employment growth, household income growth, and population growth.[9] From this perspective, in a credit expansion, credit growth would be stronger in areas with stronger tradable sector growth due to at least two reasons: (1) higher foreclosure price of house and (2) higher income growth that can be recoursed by lenders given default.

Ideal measurement of local tradable sector growth requires census-style data covering the accounting data of all firms in the tradable sector, which does not exist. Instead, I use the local employment composite (share) as a proxy for the composite of the local tradable sector. Then, I use the large time-series change in U.S. net export growth in manufacturing industries as a proxy for the relative growth (shift) at the industry level. Aggregating the above two proxies (share and shift) results in a good measure of the relative growth of the local treatable sector across metropolitan areas.[10] To account for the commuting within local areas, I aggregate net export growth at the metropolitan level.

I implement the above approach in two steps in data. First, net export measure in industry g year t is defined as $\text{NetExp}_{g,t} = \frac{Export_{g,t} - Import_{g,t}}{Y_{g,91}}$, where $Export_{g,t}$, $Import_{g,t}$, and $Y_{g,91}$ are US export, import and domestic production in industry g in year t (or 1991), respectively. All terms are measured in 2007 U.S. dollars. The scaling factor, U.S. domestic production in 1991, is chosen to avoid the potential response of domestic production to trade in later years in my study.[11]

[9] Olsen (1987) surveys on the demand factors of housing, including the above three factors.

[10] We acknowledge that the use of manufacturing does not include several other factors in the tradable sector: college town, retirement community, other tradable goods industries (e.g., natural resource), and tradable services (e.g., information technology and medical sector). Instead, I will use control variables to account for most of the above factors in Section 4.1.

[11] The same choice of domestic production in 1991 as the denominator is taken by other papers like Barrot, Loualiche,

Second, similar to Acemoglu, Autor, Dorn, Hanson, and Price (2016), I use local employment data (share) to aggregate net export growth at the metropolitan level across the period:

$$\triangle_{t_1,t_2}\text{NetExp}_m = \sum_g \left[(L_{m,g,t_0}/L_{m.t_0}) * (\text{NetExp}_{g,t_2} - \text{NetExp}_{g,t_1})\right] \quad (1)$$

where L_{m,g,t_0} and $L_{m.t_0}$ are the employment of industry g and total employment in metropolitan area m in year t_0, respectively. I choose $t_0 = t_1 - 1$ to make sure the employment share is pre-existing to the trade measure so that all changes in net export growth ($\triangle_{t_1,t_2}\text{NetExp}_m$) is completely driven by changes in trade measures (shift) rather than employment composite (share). Year t_1 to t_2 is the period of interest.

The net export growth above satisfies the requirement of the treatment variable. First, it has enough area coverage (U.S. mainland) and enough time coverage (91-09). Second, its geographic variation is large because the local tradable sector tends to cluster within several related industries due to economies of scale. Internal economies of scale increase the size of local tradable firms by at least three mechanisms: input purchase at a volume discount, fixed cost of plant operating, and learning in operation (The World Bank, 2009). Kwon, Ma, and Zimmermann (2022) recently study the corporate concentration (e.g., asset share or sales share of top-size business) by digitizing Statistics of Income (SOI) and the associated Corporate Source Book from the Internal Revenue Service (IRS) from 1918 to 2018. They find a persistent increase in the shares accounted for top-size corporations throughout the last century. External economies of scale attract firms in the same and related industries by at least three channels: specialization of suppliers, labor market pooling, and knowledge spillovers (Krugman, Obstfeld, and Melitz, 2018). Third, two other trends also strengthen the local industry clustering. The first trend is the long-term reduction in transportation costs due to the invention of new technology, construction of infrastructure, adoption of information technology, and reorganization of logistic methods (Redding and Turner, 2015). The second trend is the increased labor mobility across areas within the United States due to increased education attainment, easy availability of housing, encouraging public policies, and long-standing cultural (Greenwood, 1997; Molloy, Smith, and Wozniak, 2011). Famous examples of local industry clustering include the semiconductor industry concentrated in Silicon Valley, the automobile industry concentrated in Detroit (Holmes and Stevens, 2004), and the carpet industry concentrated in Dalton, Georgia (Krugman, 1991).

Fourth, many global events between the 1980s and 1990s promoting international trade also added to the local industry clustering at the global level. In the 1980s, many developing countries rapidly abolished earlier import substitution strategies and sometimes undertook radical trade liberalization. In addition, China's economic transition accelerated in the 1990s (Amiti and Freund, 2010), and India's reforms were initiated in 1991 (Rodrik and Subramanian, 2004). Besides, Global negotiation among the General Agreement on Tariffs and Trade (GATT) members resulted in 33

Plosser, and Sauvagnat (2022).

percent between 1980-1987, by 38 percent between 1995-1999 (World Trade Organization, 2007, p. 207). Moreover, the North American Free Trade Agreement (NAFTA) was established in 1994 (Krugman, 2008) and the World Trade Organization (WTO) in created in 1995. Further, the Dissolution of the Soviet Union in 1991 triggered many countries in Central and Eastern Europe and the former Soviet Union to apply to accede to the GATT and then the WTO (Broadman, 2006).

Fifth, the persistence feature of net export growth at the metropolitan level arises from three dimensions. At the industry level, both comparative advantages (due to technology level or natural endowment) and horizontal specialization (due to economies of scale) across nations make trade patterns persistent over time. At the local area level, clusters of manufacturing industries have been established over decades, involving extensive plants built and the migration by workers. Such industry cluster formation with huge costs is unlikely to change rapidly in a short period of time. At the individual level, human capital accumulation takes both time and investment, making job reallocation across different industries very difficult both locally and remotely.

2.2 Gravity Model-based Instrumental Variable

In this subsection, I start with the potential OLS bias and then illustrate how the instrumental variable approach can achieve causal estimates. I start with OLS specification that regresses the house price growth on the growth of private-label mortgages (non-jumbo) (PLMNJ) at the county level:

$$\triangle_{99,05} Ln(HP_c) = \beta * \triangle_{99,05} Ln(\text{Private-Mort}_c) + \gamma * Controls_c + \epsilon_c \tag{2}$$

Here, the dependent variable $\triangle_{99,05} Ln(H.P._c)$ is the house price growth (normalized to 2007 USD) at county c 99-05. The independent variable $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth of dollar amount (normalized to 2007 USD) of private-label mortgages (non-jumbo) (PLMNJ) at county c 99-05.

Potential omitted variables could bias down β. For example, the rapid net export growth and, hence, mortgage growth in 1999-2005 have been anticipated by employees in Silicon Valley so that house prices would grow before 1999 to reflect such expectations. In this case, β could be biased downward because some of the effect of the mortgage on house price shows up in house price growth in early periods, reducing the house price growth 1999-2005.

To overcome the potential endogenous concern regarding OLS specification, I use gravity model-based instrumental variable introduced by Feenstra, Ma, and Xu (2019) for net export growth.[12] They construct IV for exports and imports separately, I take a further step to combine both as an IV for net exports. To illustrate the key idea, I show the export model while keeping the import

[12] I am very grateful for help from Robert Feenstra, Hong Ma, and Yuan Xu in replicating their instrumental variables.

model in Appendix Section A.2.

As for the instrumental variable for U.S. exports at the industry-year level, Feenstra, Ma, and Xu (2019) builds on the idea that eight other high-income countries' exports can instrument for the U.S. exports since they both capture the world's rising demand. In addition, they also incorporate tariff changes, which are exogenous to foreign firms. Lastly, this method corrects for the supply shocks in the home country by using a fixed effect to remove them.

To predict U.S. export, the gravity model-based IV begins with a simple symmetric constant-elasticity equation in Romalis (2007) for export:

$$\frac{X_{s,v,t}^{US,j}}{X_{s,v,t}^{i,j}} = \left(\frac{w_{s,t}^{US} * d^{US,j} * \tau_{s,t}^{US,j}}{w_{s,t}^{i} * d^{i,j} * \tau_{s,t}^{i,j}} \right)^{1-\sigma} \tag{3}$$

In the above formula, $X_{s,v,t}^{U.S.,j}$ is U.S. export to country j in product variant v in industry s in year t. $X_{s,v,t}^{i,j}$ is the similar term but representing export from country i to j. $w_{s,t}^{US}$ and $w_{s,t}^{i}$ are the relative marginal cost of production in industry s in the US and country i. $\tau_{s,t}^{U.S.,j}$ and $\tau_{s,t}^{i,j}$ are the *ad valorem* total import tariff imposed by country j on exports from the U.S. and country i, respectively; $d^{U.S.,j}$ and $d^{i,j}$ are the bilateral distance and other fixed trade costs from U.S. to country j and from country i to country j. Lastly, σ denotes the constant elasticity of substitution ($\sigma > 1$).

The intuition of this gravity-style model is quite straightforward. Competing with country i, U.S. exports to the country j are decreasing with the ratio of relative marginal cost, the ratio of bilateral distance, and the ratio of *ad valorem* total import tariff.

Suppose that there are $N_{s,t}^{i}$ identical product varieties exported by country i to the country j in the industry s and year t, Feenstra, Ma, and Xu (2019) re-arranges the above equation, multiply both sides with $N_{s,t}^{i}$, and sum over countries $i \neq US$:

$$X_{s,v,t}^{US,j} * \sum_{i \neq US} \left[N_{s,t}^{i} (w_{s,t}^{i} d^{i,j})^{1-\sigma} \right] = (w_{s,t}^{US} d^{US,j} \tau_{s,t}^{US,j})^{1-\sigma} * \sum_{i \neq US} \left[N_{s,t}^{i} X_{s,v,t}^{i,j} (\tau_{s,t}^{i,j})^{\sigma-1} \right]$$

Since above equation holds for any countries $i \neq US$, one can choose the set of countries that have similar economic conditions (so that they are closely competing with U.S. exports) to make my prediction more accurate. Following Autor, Dorn, and Hanson (2013b), Feenstra, Ma, and Xu (2019) use the eight high-income countries.

Again, I can multiple $N_{s,t}^{U.S.}$ (number of variants of products exported by U.S.) on both sides and denote the sectoral exports $X_{s,t}^{U.S.,j} \equiv X_{s,v,t}^{U.S.,j} N_{s,t}^{U.S.}$ and $X_{s,t}^{i,j} \equiv X_{s,v,t}^{i,j} N_{s,t}^{i}$, then I have

$$X_{s,t}^{US,j} * \sum_{i \neq US} \left[N_{s,t}^{i} * (w_{s,t}^{i} d^{i,j})^{1-\sigma} \right] = N_{s,t}^{US} * (w_{s,t}^{US} d^{US,j} \tau_{s,t}^{US,j})^{1-\sigma} * \sum_{i \neq US} \left[X_{s,t}^{i,j} * (\tau_{s,t}^{i,j})^{\sigma-1} \right]$$

After a few re-arrangements, one can get the formula for $X_{s,t}^{U.S.,j}$:

$$X_{s,t}^{US,j} = \frac{N_{s,t}^{i} * (w_{s,t}^{US} d^{US,j} \tau_{s,t}^{US,j})^{1-\sigma}}{\sum_{i \neq US} \left[N_{s,t}^{i} * (w_{s,t}^{i} d^{i,j})^{1-\sigma} \right]} * \left(\sum_{k \neq US} X_{s,t}^{k,j} \right) * \sum_{i \neq US} \left[\frac{X_{s,t}^{i,j}}{\sum_{k \neq US} X_{s,t}^{k,j}} * (\tau_{s,t}^{i,j})^{\sigma-1} \right] \tag{4}$$

Note in the above formula, I multiply and divide by $\sum_{k \neq U.S.} X_{s,t}^{k,j}$ to prepare for the regression setup in the next step. Now I can take logs of the above equation and get the regression-style formula:

$$\begin{aligned} \ln\left(X_{s,t}^{US,j}\right) = & \underbrace{\ln\left(\sum_{k \neq US} X_{s,t}^{k,j} \right)}_{\text{Term 0}} + \underbrace{\ln\left(N_{s,t}^{US} (w_{s,t}^{US})^{1-\sigma}\right)}_{\text{Ind-Year FE: } \alpha_{s,t}^{US}} + \underbrace{(1-\sigma)\ln\left(d^{US,j}\right)}_{\text{Importing-country FE: } \delta^{US,j}} \\ & + \underbrace{(1-\sigma)\ln\left(\tau_{s,t}^{US,j}\right)}_{\text{Term 1}} + \underbrace{(\sigma-1)\ln\left(\left\{ \sum_{i \neq US} \left[\frac{X_{s,t}^{i,j}}{\sum_{k \neq US} X_{s,t}^{k,j}} (\tau_{s,t}^{i,j})^{\sigma-1} \right] \right\}^{\frac{1}{\sigma-1}} \right)}_{\text{Term 2:}(\sigma-1)\ln\left(T_{s,t}^{j}\right)} + \epsilon_{s,t}^{j} \end{aligned} \tag{5}$$

We can see that U.S. exports to the country j in the industry s year t can be decomposed into six terms. "Term 0" is the exports from eight other high-income countries (whose economic development is similar to the U.S.) to the destination country j, which reflects the world demand. The second term $\alpha_{s,t}^{U.S.}$, which reflects the U.S. supply shocks by industry and year, is potentially endogenous. For instance, the U.S. computer chip industry exports a lot, partially because of its superior investment in human capital and innovation. Feenstra, Ma, and Xu (2019) remove this term by the U.S. industry-year fixed effects. The third term $\delta^{U.S.,j}$ reflects the distance from the U.S. to the destination market j and all other industry- and year-invariant trade costs. For example, the US exports more to Mexico than the mid-east, partially because of the lower transportation cost associated with shorter distances. Since this term is predetermined and endogenous to firms' exporting decisions, Feenstra, Ma, and Xu (2019) remove it by importing-country fixed effects. "Term 1" represents the tariffs on U.S. exports imposed by destination country j, which is out of control by U.S. firms. This term is retained to capture the shocks from tariffs. "Term 2" is the weighted average tariffs on non-US exports imposed by destination country j, which is again out of control of U.S. firms. Intuitively, when this weighted average tariff on non-US exports rises, destination country j will import more from the U.S. as substitutions. This term is retained to reflect such substitution effect. The last term $\epsilon_{s,t}^{j} = -\ln\left(\sum_{k \neq US} [N_{s,t}^{i} (w_{s,t}^{i} d^{i,j})^{1-\sigma}]\right)$ is unobserved and remains in the regression error term.

After the above regression, I can construct predicted U.S. exports that are isolated from supply shocks and the predetermined factor by the above "Term 0", "Term 1", and "Term 2":

$$\ln\left(\widehat{X_{s,t}^{US,j}}\right) = \ln\left(\sum_{k \neq US} X_{s,t}^{k,j} \right) + \hat{\beta}_1 * \ln\left(\tau_{s,t}^{US,j}\right) + \hat{\beta}_2 * \ln\left(T_{s,t}^{j}\right) \tag{6}$$

Data Implementation There are four detailed steps implementing the above procedures in data

to get net export and its GIV at the metropolitan level across periods. First, I estimate Eq (5) at the 6-digit H.S. industry level (5673 industries) and get predicted U.S. export (isolated from supply shocks) by Eq (6). Second, I aggregate predicted U.S. exports across importing countries and crosswalk 6-digit H.S. code (5673 industries) to 4-digit revised SIC system (392 manufacturing industries) via the crosswalk with weights by Acemoglu, Autor, Dorn, Hanson, and Price (2016). I end up with predicted U.S. exports to the world at the industry g year t level. We get this aggregation as $\widehat{X_{g,t}^{US}} = \sum_{s \in g} \sum_j \widehat{X_{s,t}^{US,j}}$. I do a similar work to get predicted U.S. imports from the world $\widehat{M_{g,t}^{U.S.}}$. Third, I derive the gravity model-based instrumental variable for net export at the industry-by-year level

$$\text{givNetExp}_{g,t}^{US} = \frac{\widehat{X_{g,t}^{US}} - \widehat{M_{g,t}^{US}}}{Y_{g,91}} \tag{7}$$

where $Y_{g,91}$ represents US domestic production in year 1991. Fourth, similar to my construction in net export measure, I use employment data from County Business Pattern to aggregate $givNEP$ at the metropolitan area across periods.

$$\triangle_{t_1,t_2}\text{givNetExp}_m = \sum_g \left[(L_{m,g,t_0'}/L_{m.t_0'}) * (\text{givNetExp}_{g,t_2} - \text{givNetExp}_{g,t_1}) \right] \tag{8}$$

where $L_{m,g,t_0'}$ and $L_{m.t_0'}$ are the employment of industry g and total employment, respectively, in metropolitan area m in year t_0'. Following Acemoglu, Autor, Dorn, Hanson, and Price (2016), I choose $t_0' = t_1 - 3$ to avoid potential data error covariance between dependent variable and independent variable.

Relevance Condition Intuitively, the gravity model-based IV captures the exogenous part of net export growth due to (1) rising world demand reflected in net export growth by other eight high-income counties and (2) tariff changes, after controlling for U.S. industry-by-year supply shocks. This relevance condition of this GIV is satisfied since it is derived from the general equilibrium model by Romalis (2007) and specific decomposition by Feenstra, Ma, and Xu (2019). I will test this condition in data by robust F-statistics (Kleibergen and Paap, 2006) and efficient F-statistics Olea and Pflueger (2013).

Exclusion Restriction There are exclusion restrictions at two levels. The first level refers to the gravity model-based instruments by Feenstra, Ma, and Xu (2019). They have removed supply-side shocks by industry fixed-effect in predicted U.S. exports and demand-side shocks by industry fixed-effect in predicted U.S. imports. Thus, exclusion restriction holds for the gravity model-based IV. The second level refers to my regression specification Eq (2), in which gravity model-based IV serves as the IV for the growth in private-label mortgages (non-jumbo). Exclusion restriction here requires that net export growth impacts house prices only through its role in contributing to the growth of private-label mortgages (non-jumbo). Empirically, the IV estimation of the impact of private-label mortgages (non-jumbo) on house prices does not separate demand and supply channels brought about by exogenous changes in net exports. Instead, I rely on additional tests to address

the concern of demand-side (speculation) channels in 5.

As with all instrumental variable estimates, my 2SLS estimates represent the local average treatment effects on compliers (Imbens and Angrist, 1994). In my setting, compliers are metropolitan-by-year observations that experience more industry-by-year U.S. net export to the world following increases in gravity model-predicted U.S. net export.

3 Data Sources

I combine several data sources to study how credit expansion causes a stronger housing boom and bust in the high-net-export-growth areas (HNEG areas). The granular international trade and tariff data are new to the literature on housing, allowing the use of the gravity model-based instrumental variable from international economics (Feenstra, Ma, and Xu, 2019) as my identification strategy.

3.1 Data for Net Export Growth

Trade Flow Data International trade flow data for 1991–2011 are from the UN Comtrade Database.[13] This database records bilateral imports and exports data for detailed products recorded under the six-digit Harmonized Commodity Description and Coding System (HS code). For the gravity model-based IV, I use trade data for ten countries: USA, China, and the eight other high-income countries (Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland) used by Autor, Dorn, and Hanson (2013a). To deflate trade value to 2007 USD dollar, I use the Personal Consumption Expenditures Chain-type Price Index from Federal Reserve in St. Louis.[14]. To crosswalk these trade data from a six-digit HS code to a four-digit SIC code, I use the crosswalk file and revised SIC system (392 manufacturing industries) in Acemoglu, Autor, Dorn, Hanson, and Price (2016).[15] [16]

Tariff Data I obtain Bilateral tariff schedule data at five-digit SITC products between 1984 to 2011 from Feenstra and Romalis (2014).[17] To crosswalk tariff data from a five-digit SITC system

[13] UN Comtrade Database is available here: `https://comtrade.un.org/data/`.

[14] Personal Consumption Expenditures Chain-type Price Index from in St. Louis Federal Reserve is available at `https://fred.stlouisfed.org/series/PCEPI`

[15] This crosswalk file is also available from Prof. David Dorn's website: `https://www.ddorn.net/data.htm`. The file name is "[D4] HS 1996/2002 6-digit to sic87dd 4-digit". The further refined SIC system (392 manufacturing industries) and crosswalk file are available from Acemoglu, Autor, Dorn, Hanson, and Price (2016).

[16] To make sure my calculation is correct, I calculate China's exports to the US and eight other high-income countries at the industry-year level from 1991 to 2007 and compare them to data provided by David Dorn's website. The trade data is within the section [D] Industry Trade Exposure via his data page `https://www.ddorn.net/data.htm`. Correlations between my calculation and his corresponding data are 0.9983 for China's export to the U.S. and 0.9973 for China's export to eight other high-income countries.

[17] The original data are collected from the TRAINS and IDB databases accessed via the World Bank's WITS website and various other resources, with multiple steps of cleaning and replacing missing values by other resources. The World Bank's WITS website is `https://wits.worldbank.org/`. The complete procedure of data work is described in Appendix C in Feenstra and Romalis (2014).

to a six-digit HS system, I follow the methods in Feenstra, Ma, and Xu (2019). Specifically, I first convert the HS 2007 version to the HS 2002 version by the crosswalk files from the Trade Statistics Branch (TSB) of the United Nations Statistics Division.[18] Then I employ a crosswalk from Feenstra, Lipsey, Deng, Ma, and Mo (2005) to match each six-digit HS code to one 5-digit SITC2. In the cases that one six-digit HS code is matched to multiple SITC2 codes, I follow Feenstra, Ma, and Xu (2019) and employ the one with the highest value share.

Manufacturing Production Data I obtain the U.S. 4-digit SIC manufacturing industry total domestic production (vship) in 1991 from the NBER-Center for Economics Studies (NBER-CES) Manufacturing Industry Database. I use the industry total domestic production in the first year 1991 in my sample as the denominator to scale the trade value so the trade change afterward cannot respond to it.[19]

Manufacturing Employment Data I obtain Employment data in detailed manufacturing industries at the county-by-year level from County Business Patterns Database at U.S. Census.[20] Following Acemoglu, Autor, Dorn, Hanson, and Price (2016), I use manufacturing employment data to aggregate net export growth and its IV at the metropolitan areas across periods. For details, refer to Section 2.2.

3.2 Data for Mortgages and House Prices

Mortgage Data Detailed loan-level mortgage data are from the Home Mortgage Disclosure Act (HMDA) database.[21] In 1975, Congress enacted HMDA to improve public access to information of mortgage loans. Any financial institution must report HMDA data to its designated regulator if it meets certain criteria, such as a threshold for asset size and if the institution has a home office or branch in a Metropolitan Statistical Area (MSA).[22] This annual database contains information on loan applications (regardless of whether they are approved or denied), borrower demographics, lender identifiers, and loan specifics such as purpose, amount, and location. The HMDA database provides near-universal coverage of the mortgage market. Avery, Bhutta, Brevoort, Canner, Gibbs,

[18]The crosswalks files between different HS versions are available from the UN Comtrade database: `https://unstats.un.org/unsd/trade/classifications/correspondence%2Dtables.asp`.

[19]This choice of scaling is also used by Barrot, Loualiche, Plosser, and Sauvagnat (2022).

[20]County Business Patterns Database is available here: `https://www.census.gov/programs-surveys/cbp/data/datasets.html`.

[21]The Federal Financial Institutions Examination Council (FFIEC) releases 2017-2021 HMDA data at `https://ffiec.cfpb.gov/data%2Dpublication/2021`. The Consumer Financial Protection Bureau (CFPB) releases 2007-2017 HMDA data at `https://www.consumerfinance.gov/data%2Dresearch/hmda/historic%2Ddata/`. CFPB provides links to 1990-2006 HMDA contained in National Archives at `https://github.com/cfpb/HMDA_Data_Science_Kit/blob/master/hmda_data_links.md`. Alternatively, the University of Michigan provides 1990-2006 HMDA data at `https://www.openicpsr.org/openicpsr/project/151921/version/V1/view`.

[22]For a depository institution that has a home office or a branch in an MSA, it must file HMDA data if it has made a home purchase loan or has refinanced a home purchase loan and has assets above a threshold set up annually by the CFPB. For a non-depository institution (e.g., a mortgage company that raises funds instead from banks or capital markets), it must file HMDA data if at least 10% of the loan portfolio is comprised of home purchase loans and if its asset is above $ 10 million.

et al. (2010) conclude that commercial banks filing HMDA in 2008 carried 93% of the total mortgage dollars outstanding on commercial bank portfolios.[23]

I use the following filtering criteria for the raw HMDA data. First, I keep originated loans and drop applications that are denied, withdrawn, or not accepted for other reasons. Second, for loan types, I keep conventional and Federal Housing Administration-insured (FHA-insured) loans and drop mortgages insured by the Veterans Administration, Farm Service Agency, or Rural Housing Service. Third, for loan purposes, I use home purchase mortgages for most empirical tests and include refinancing mortgages for some robustness tests. Fourth, for occupancy types, we keep both owner-occupied and non-owner-occupied mortgages and treat "not applicable" occupancy as owner-occupied.[24]

I use the HMDA database to construct loan volume (number and dollar amount) at the county-by-year level for private-label mortgages (PLMs) and government-sponsored enterprise mortgages (GSEMs). According to the HMDA examination procedures, an institution must report the type of entities that purchase the mortgages that are originated (or purchased) and then sold in the same calendar year.[25] I use the classification of PLMs and GSEMs by Mian and Sufi (2022).[26]

Conforming Loan Limits Data I obtain conforming loan limits by year and county from Federal Housing Finance Agency.[27] Conforming loan limits are, in general, different for dwellings with different units in each year (and county after 2007). Since the HMDA data does not contain information on the number of units in a home between 1991 and 2009, I use the 1-unit conforming loan limit for all mortgages. Thus my measure of non-jumbo mortgages is potentially an underestimation of the true size, avoiding an upward bias of my results.

Consistent Counties Covered by HMDA Although HMDA only mandates mortgage reports in metropolitan areas, many institutions still report all mortgages regardless of location or requirement on institutions. Thus, from the HMDA dataset, I cannot infer the required counties based on non-missing data. Instead, I need to find counties that are consistently covered by HMDA based on

[23] Even though lenders with offices only in non-metropolitan areas are exempt from filing HMDA, 83.2% of the population in 2006 lived in metropolitan areas (Dell'Ariccia, Igan, and Laeven, 2012).

[24] According to the HMDA manual (`https://www.ffiec.gov/hmda/pdf/1998guide.pdf`), this "not applicable" occupancy very likely refers to a multifamily dwelling the borrower lives in. In terms of magnitude, this "not applicable" occupancy is only around 0.59% of the number of owner-occupied loans and 3.5% of the number of non-owner-occupied loans as of 2007.

[25] The "Home Mortgage Disclosure Act Examination Procedures" is available at `https://www.federalreserve.gov/boarddocs/caletters/2009/0910/09%2D10_attachment.pdf`. These procedures imply that the HMDA can potentially under-estimate the mortgages that are sold as GSEMs and PLMNJs since the mortgages originated near the end of a calendar year need some time for sale. Nonetheless, this potential underestimation can only bias my results to zero.

[26] Mian and Sufi (2022) group five categories as PLMs when a mortgage is sold: (1) into private securitization, (2) to a commercial bank, savings bank, or savings affiliation affiliate, (3) to a life insurance company, credit union, mortgage bank, or finance company, (4) to an affiliate institution, and (5) to other types of purchaser.

[27] Before and in 2007, FHFA sets conforming loan limits only at the national level: `https://www.fhfa.gov/AboutUs/Policies/Documents/Conforming%2DLoan%2DLimits/loanlimitshistory07.pdf`. From 2008 onward, FHFA sets conforming loan limits by year and by county: `https://www.fhfa.gov/DataTools/Downloads/Pages/Conforming%2DLoan%2DLimit.aspx`.

the scope of metropolitan areas defined and updated over time by the U.S. Office of Management and Budget. Avery, Brevoort, and Canner (2007) point out that metropolitan boundaries were reasonably stable from 1996 to 2003 but changed a lot in 2004 HMDA due to the updated 2003 version of metropolitan areas. Thus, I define "HMDA consistent counties after 1996" as counties that are covered in the following two versions of the metropolitan definition: (1) metropolitan areas definition (MSA and CMSA code) of 1999 version[28] and (2) metropolitan areas definition (CBSA code) of 2003 version[29]. In addition, I require number of loans is positive in the initial year (1996) for three mortgage categories: retained in the balance sheet, GSEM and PLMNJ. Consequently, "HMDA consistent counties after 1996" contains 800 counties. Likewise, for the sample period that starts from 1991 or 1992, I define "HMDA consistent counties after 1990" as counties that are covered in the following two sets of metropolitan areas: (1) metropolitan areas definition (MSA and CMSA code) of 1990 version[30] and (2) the above "HMDA consistent counties after 1996". Again, I require the number of loans in 1991 is positive for three mortgage categories: retained in the balance sheet, GSEM and PLMNJ. Consequently, "HMDA consistent counties after 1990" contains 712 counties.

U.S. House Price Data The U.S. house price index data based on repeated sales at the county and ZIP levels are from the Federal Housing Finance Agency.[31] This database has reasonably good coverage of the US counties in metropolitan areas in and after 1991 (870 out of 1084 counties in metropolitan areas in 1991).

Merge House Price and Mortgage Data For both figure and regression analysis regarding house prices, I require that the counties are covered by both house price data and mortgage data. The merged data set contains fewer counties compared to mortgage data since house price data covers fewer counties. Second, I require that each county contains both data in the period 1991-2011 (1999-2011) for the figure (regression) analysis. For more data details regarding mortgage and house prices at the ZIP level, refer to Section A.1.

3.3 Local Economic Conditions

BEA Employment Data To test the impact of net export growth on total employment, I obtain annual employment data at the county level from the U.S. Bureau of Economic Analysis (BEA)

[28] Metropolitan areas definition (MSA and CMSA code) of 1999 can be found here: `https://www2.census.gov/programs%2Dsurveys/metro%2Dmicro/geographies/reference%2Dfiles/1999/historical%2Ddelineation%2Dfiles/99mfips.txt`. At the moment of data work, the above link does not work. I use crosswalk from the county to the MSA in 2000 from Missouri Census Data Center at `https://mcdc.missouri.edu/applications/geocorr2000html`.

[29] Metropolitan areas definition (CBSA code) of 2003 can be found here: `https://www2.census.gov/programs%2Dsurveys/metro%2Dmicro/geographies/reference%2Dfiles/2003/historical%2Ddelineation%2Dfiles/030606omb%2Dnecta%2Dcnecta.xls`.

[30] I use crosswalk from the county to the MSA in 1990 from Missouri Census Data Center at `https://mcdc.missouri.edu/applications/geocorr1990.html`.

[31] The data are available at `https://www.fhfa.gov/DataTools/Downloads`.FHFA working paper Bogin, Doerner, and Larson (2019) describes the construction of the index and tests its accuracy via various methods.

because such data include both (1) wage and salary employment and (2) proprietor employment (self-employment).[32] This coverage is better than County Business Pattern employment data, which does not include self-employment outside of establishments.

IRS Household Income Data To test the impact of net export growth on household income growth, I use the annual data at the county level from the Statistics of Income Division in the U.S. Internal Revenue Service (IRS).[33] This income data is based on the addresses reported on individual income tax returns filed with the IRS. I calculate average household income at the county level as the adjusted gross income divided by the number of returns (households). Then the income is adjusted to the 2007 USD by the Personal Consumption Expenditures Chain-type Price Index (PCEPI) from the Federal Reserve Bank of St. Louis.

Working-Age Population Data In testing the impact of net export growth on the local population growth, I obtain the annual data on population at the county level from the U.S. Census.[34]

Local Control Variables Control variables at county and ZIP code levels are from U.S. Decennial Census Summary Files. For the year 1989, control variables at the county level are from 1990 (March) Census Summary File 1C and 3C, while control variables at the ZIP level are from Summary File 3B.[35] For the year 1999, control variables at the county level and the ZIP level are from both 2000 (March) Census Summary File 1 and 3.[36]

3.4 Twelve Counties Severely Affected by 2005 Hurricanes

Twelve Counties Severely Affected by 2005 Hurricanes I delete twelve "deeply affected counties by 2005 Hurricanes" since they experienced unusual growth in mortgages due to hurricane damage and subsequent government subsidies.[37] In 2005, three Category 5 hurricanes (Katrina, Rita, and Wilma) caused enormous fatalities and damage in Alabama, Florida, Louisiana, Mississippi, and Texas. Among them, Katrina was the most devastating natural disaster in United States history, causing 1392 fatalities, 200,000 homes destroyed (Deryugina, Kawano, and Levitt, 2018), and an estimated $125 billion in damage in 2005 US dollars. The deeply affected counties have had unusual growth in mortgages in 2006. First, many households may choose to stay away

[32]The BEA employment data is available at `https://apps.bea.gov/regional/downloadzip.cfm`. Under the category "Personal Income (State and Local)", "CAEMP25S" contains data from 1969 to 2000, and "CAEMP25N" contains data in and after 2001.

[33]For 1989 to 2018, the data is available at `https://www.irs.gov/statistics/soi%2Dtax%2Dstats%2Dcounty%2Ddata`.

[34]The data is available here: https://www.census.gov/programs-surveys/popest.html

[35]The 1990 U.S. Decennial Census File 1 database is available at `https://www.census.gov/data/datasets/1990/dec/summary-file-1.html` and File 3 database is available at `https://www.census.gov/data/datasets/1990/dec/summary-file-3.html`.

[36]The 2000 U.S. Decennial Census Files are available here: `https://www.census.gov/programs-surveys/decennial-census/guidance/2000.html`

[37]I try my best to present the most robust results. Since outliers only largely affect results in regression but not the illustration in figures, I include these twelve counties in the figures but remove them from regressions and summary statistics.

permanently from the disaster areas. Second, lots of households may delay mortgage applications due to financial constraints caused by such hurricanes. Third, government assistance programs may replace much of the role played by PLM (and even GSEM). For example, St. Bernard Parish in Louisiana (county FIPS code 22087) experienced a negative 66.6% change in PLMNJ between 2002 to 2006, mainly due to the Katrina hurricane. For the above reasons, I delete twelve "deeply affected counties by 2005 Hurricanes". I define "deeply affected" counties when more than 10 percent of occupied units were identified as major or severe damage by the Federal Emergency Management Agency as of February 2006.[38]. I use the percentage of major and severe damage since these damages are more likely to be reported by households and inspected by federal agencies, whereas minor damage might be under-reported. Eventually, I drop a list of the twelve most-affected counties.[39]

3.5 Summary Statistics and Figures

Summary statistics of key variables are reported in Table (1), separated into different periods. In the prior period (1991-1999), starting from 712 "HMDA consistent counties after 1990", we delete seven counties due to severe impact of the 2005 hurricanes described in section (3.2), resulting in 705 counties. Due to data availability, we have fewer observations for house price growth and housing supply elasticity. In the boom period (1999-2005) and bust period (2007-2009), starting from 800 "HMDA consistent counties after 1996", we delete eight counties due to severe impact of the 2005 hurricanes, resulting in 792 counties.

First, we can see that the house price experienced a clear and pronounced cycle. While the mean of annualized growth rate across counties is only 1.30% in the prior period (1991-1999), it is 3.82% in the boom period (1999-2005) and -5.4% in the bust period (2007-2009).

Second, the government-sponsored enterprise mortgages (GSEM) play an less important role in the boom period: its mean annualized growth rate across counties is 17.07% in the prior period but only 4.38% in the boom period.

Third, for the private-label mortgages (non-jumbo) (PLMNJ), it is interesting that the mean of growth rates across counties is similar between the prior period (17.07%) and boom period (16.98%). The similar mean values result from the use of different base values.[40] The similar mean growth rates mask the much larger increase in absolute dollar amount in the boom period in private-label mortgages. Figure (11) shows the differential increase in dollar amount in PLMNJ

[38]The Federal Emergency Management Agency reports the estimates of housing unit damage on Pages 15-19 of the following report at `https://www.huduser.gov/publications/pdf/gulfcoast_hsngdmgest.pdf`.

[39]These most-affected counties include Monroe County (FL, 12087), Cameron Parish (LA, 22023), Jefferson Parish (LA, 22051), Orleans Parish (LA, 22071), Plaquemines Parish (LA, 22075), St. Bernard Parish (LA, 22087), St. Tammany Parish (LA, 22103), Vermilion Parish (LA, 22113), Hancock County (MS, 28045), Harrison County (MS, 28047), Jackson County (MS, 28059), Stone County (MS, 28131).

[40]To be specific, the 1991 PLMNJ dollar amount is used for the growth rate between 1991 and 1999 whereas the 1999 PLMNJ dollar amount is used for the growth rate between 1999 and 2005.

in the boom period, because time series values are scaled by 1991 dollar value only. Figure (11) and (10) hint that the impact of private-label mortgages (non-jumbo) will be much larger than the government-sponsored enterprise mortgages in the boom period since the former ones do show a differentially larger increase in the high-net-export metropolitan areas.

Fourth, the net export growth rates are a little bit more negative in the boom period (with a mean of -0.22% and standard deviation of 0.20%) and in the prior period (with a mean of -0.18% and standard deviation of 0.17%). This data pattern is also present in the gravity model-based net export growth rates.

We include control variables at county c only in the starting year in each period, which prevents any impact from net export growth during the prior period. Control variables are used to neutralize factors that may affect credit expansion for reasons unrelated to the main hypothesis. Basic controls include the number of households, average household income, and the fraction of the labor force at the county c. Housing controls include the number of house units, housing supply elasticity (Saiz, 2010), house vacancy rate, and fraction of renters in the occupied house units. My demographic controls include the fraction of the population holding a Bachelor's degree and above, the percentage of the white population, and the count of immigrants entering the U.S. between 1990 and 1999. Industry controls include the ratio of the population (1) in the art, entertainment, and recreation industries, (2) in the health industries, (3) in the tradable service industries, and (2) as college students. The industry controls capture the phenomena of retirement towns, medical centers, and college towns and the effect of the tradable service sector.

4 Empirical Results: Supporting Credit Expansion

My main empirical tests in this section provide direct evidence that credit expansion in private-label mortgages (PLMs) rather than government-sponsored enterprise mortgages (GSEMs) cause the 1999-2009 housing price boom and bust. To compare these two types of mortgages under the same criteria, I use the conforming loan limits to get the non-jumbo category of private-label mortgage (PLMNJ) and remove the jumbo ones. I will focus on the comparison between private-label mortgages (non-jumbo) (PLMNJs) and the government-sponsored enterprise mortgages (GSEMs).[41]

Based on the intuition of the model described in the introduction, I conduct three tests for my set of results. My first test shows that, induced by net export growth, credit expansion in (non-jumbo) private-label mortgages (PLMNJs) cause the 1999-2009 housing price boom and bust across metropolitan areas in the USA. My second test shows that government-sponsored enterprise mortgages (GSEMs) do not respond to net export growth, and thus are unrelated to differential

[41] Jumbo loans of private-label mortgages are smaller in number and dollar amount when compared to non-jumbo loans. Including jumbo ones only strengthens my results.

house price booms and busts across metropolitan areas. The second test can also help argue against the demand (or speculation) view that mortgage growth can be driven by demand (or speculation) induced by net export growth. Using the fact that there is no aggregate credit expansion during 1991-1999, my third test provides evidence supporting the exclusion restriction of my IV strategy: net export growth impacts house price growth only through private-label mortgages.

4.1 House Price Boom (99-05) and Bust (07-09)

Let us first focus on my main hypothesis: house prices as a result of credit expansion in response to net export growth across metropolitan areas. First, I expect private-label mortgages (non-jumbo) would experience stronger growth in the high-net-export-growth metropolitan areas. Second, I expect that the drop in house prices will also be stronger in the high-net-export-growth metropolitan areas between 2007 and 2009.[42]

Hypothesis 1 In cross-section, growth in private-label mortgages (non-jumbo) (PLMNJs) in the boom period (1999-2005) causes the house price boom (1999-2005) and bust (2007-2009).

House Price Growth in Boom (99-05) To test house price boom, I perform regression with the following specifications.

$$\triangle_{99,05} Ln(HPI_c) = \beta * \triangle_{99,05} Ln(PLMNJ_c) + \gamma * \boldsymbol{Controls_c} + \epsilon_c \tag{9}$$

The left-hand-side dependent variable $\triangle_{99,05} Ln(HPI_c)$ is the growth rate of the house price index at county c 99-05, and the key variable of interest $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgage (non-jumbo) (PLMNJ) at county c 99-05. $Controls_c$ indicates control variables at county c in 1999. I use the gravity model-based instrument ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05} Ln(PLMNJ_c)$.[43]

Each regression is weighted by the natural logarithm of the number of house units in 1999. Logarithm rather than the absolute number of house units is chosen to prevent results from being dominated by a few super-populous counties. To take into account that households might commute to work across counties within a metropolitan area, I measure net export growth in the metropolitan area.[44] Furthermore, I cluster standard errors at the metropolitan area (MA) level.

Table (3) reports OLS, reduced-form, second-stage, and first-stage results in panel A, B, C, and D, respectively. First, panel A column (1) shows the positive and significant impact of PLMNJ growth (99-05, annualized) on house price index growth (99-05, annualized) without any control.

[42] The Great Recession period, 2007-2009, is defined by the US National Bureau of Economic Research at `https://www.nber.org/research/data/us-business-cycle-expansions-and-contractions`

[43] Compared to the HMDA sample starting from 1991, I have an increased sample size here due to the larger coverage of metropolitan definition after 1996. Another slight difference in the sample size arises from the data availability in house prices.

[44] According to US Census, "the general concept of a metropolitan statistical area is that of a core area containing a substantial population nucleus, together with adjacent communities having a high degree of economic and social integration with that core." (`https://www.census.gov/programs-surveys/metro-micro/about.html`)

Through columns (2) to (4), the estimates show that the OLS coefficients of PLMNJ are of the similar magnitude as more controls are added . Panel B reports the reduced-form estimates that share a similar pattern. The first-stage estimates in panel D are positive and significant.

Panel C reports 2SLS estimates for equation (9). Like reduced-form estimates, the 2SLS estimates are statistically significant. As for the weak IV concern, clustered Kleibergen-Paap F-statistic (Kleibergen and Paap, 2006) is 11.22, and the Montiel Olea-Pflueger Efficient F-Statistic (Olea and Pflueger, 2013) is 11.20, both of which are larger than 10. Thus, my estimates are very unlikely biased by a weak instrumental variable. Panel C reports 2SLS estimates for equation (9). Like reduced-form estimates, the 2SLS estimates are statistically significant at a one percent level and quite stable across various specifications. As emphasized by Jiang (2017), it is important to compare the 2SLS estimates with OLS estimates to alleviate the "blown-up" concern that statistical significance may pick up a weak IV that biases the estimate toward a large magnitude. Besides two large F-Statistics, I also confirm that 2SLS estimates are only about three to four times of the OLS estimates in my baseline specification in column (5). Thus, my estimates are very unlikely to be subject to the "blow-up" concern of a weak IV.

Let us focus on economic meaning in 2SLS with controls in Table (2). According to column (6) with all controls, one standard deviation in cross-sectional difference in annualized PLMNJ growth results in $8.05\% \times 0.438 = 3.54\%$ difference in annualized house price index (HPI) growth across metropolitan areas, translating into 21.16% difference from 1999 to 2005. One standard deviation in cross-section difference in annualized HPI growth is 3.42%, translating into 20.53% from 1999 to 2005. The two results mean that, from 1999 to 2005, one standard deviation in PLMNJ growth can explain $21.16\%/20.53\% = 103.08\%$ of one standard deviation in HPI growth.

Similarly, one standard deviation in cross-section housing supply elasticity (HSE) can explain $1.22 * 0.005 = 0.61\%$ difference in annualized HPI growth, translating into 3.65% difference from 1999 to 2005. This result implies that, from 1999 to 2005, one standard deviation in HSE can explain $3.65\%/20.52\% = 17.79\%$ of one standard deviation in HPI growth. Therefore, growth in private-label mortgages (non-jumbo) exhibits five times the explanatory power as housing supply elasticity (Saiz, 2010).

House Price Boom (99-05) and Bust (07-09) We then compare the house price boom and bust in a single regression test. My stacked 2SLS regression is

$$\begin{aligned}\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = \beta_{99,05} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{07,09} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} \\ + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{07,09} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_{period,c}\end{aligned} \tag{10}$$

Controls, weight, and standard errors are the same as Eq(9).

Table (5) reports OLS, reduced-form, second stage, and the first stage of the stacked regression of house price index in the boom period (99-05) and the bust period (07-09) in panel A, B, C,

and D.[45] First, panel A shows that the OLS coefficients of PLMNJ growth (99-05) are positive in the boom period (99-05) but negative in the bust period (07-09). The same trend applies to reduced-form estimates in panel B and 2SLS estimates in panel C. Importantly, the coefficient magnitudes in 2SLS are between three to four times of OLS results in baseline specification in column (5), much lower than the average nine in the top three finance journals. Same as the Table (3), first-stage estimates in panel D show the stable and strong positive correlation between the PLMNJ and gravity model based IV for Net Export Growth (GIV-NEG), with large enough first-stage F-Statistic. Therefore, I do not have weak IV concerns. Let us turn our attention to the coefficient equality test of PLMNJ growth in the boom and bust periods in 2SLS in the Table (4). For all specifications through columns (1)-(5), the chi-square statistics are large, and p-values are below 0.01, meaning the two coefficients are statistically different.

In terms of economic meaning, one standard deviation in cross-sectional difference in annualized PLMNJ growth (99-05) results in $8.05\% \times -0.896 = -7.21\%$ difference in the annualized house price drop (07-09). Since the lengths of the boom and bust periods are different, I need to consider the time horizon: a longer period of credit expansion results in a shorter period of bust period. One standard deviation in cross-sectional difference in six-year PLMNJ growth (99-05) can result in $7.21\% \times 2 = 14.43\%$ cross-sectional difference in the house price drop. For annualized HPI drop, one standard deviation in cross-section difference is 6.15%, translating into $6.15\% \times 2 = 12.31\%$ from 2007 to 2009. The two results mean that one standard deviation in cross-sectional difference in six-year PLMNJ growth (99-05) can explain $14.43\%/12.31\% = 117.21\%$ of one standard deviation in two-year HPI drop (07-09). To sum up, six-year PLMNJ growth can explain 103.08% house price growth 1999-2005 and 117.21% house price drop 2007-2009.

However, my specification in column (5) shows that the housing supply elasticity cannot impact house price drop after controlling all other factors. In terms of economic magnitude, one standard deviation in cross-section housing supply elasticity (HSE) can explain $1.217 * 0.04722\% = 0.057\%$ difference in annualized HPI drop (07-09), translating into 0.115% difference from 2007 to 2009. This figure implies that one standard deviation in HSE can explain $0.115\%/12.310\% = 0.933\%$ of one standard deviation in house price drop from 2007 to 2009. This explanatory ratio of 0.933% in the bust period 2007-2009 is much smaller than the ratio of 17.79% in the boom period 1999-2005.

A comparison of the explanatory ratio of private-label mortgages (non-jumbo) and housing supply elasticity, together with my comparison of growth rates in private-label mortgage (non-jumbo) and the government-sponsored enterprise mortgages in the boom period in Section (5.1), hints that the default in private-label mortgage in the bust period likely dominates the house price drop. Future research can use detailed loan-level performance data to confirm that the high-net-

[45] Please note that I require counties to be consistently covered by HMDA and HPI databases since 1999. Because of the larger inclusion of metropolitan areas in 1999 than in 1991, my sample size is increased to $786 \times 2 = 1572$ counties in columns (1)-(4) and $695 \times 2 = 1390$ in column (5) due to data availability of housing supply elasticity.

export-growth areas suffer a much larger number of defaults in private-label mortgages during the bust period (2007-2009) than the low-net-export-growth areas.

In summary, empirical tests in this subsection support the main conclusion that credit expansion in private-label mortgages induced by net export growth causes the 1999-2009 U.S. house price boom and bust across metropolitan areas. Further, in Section A.4, I will show that this main conclusion is robust to state-level differences in anti-predatory lending laws (Di Maggio and Kermani, 2017), recourse laws Ghent and Kudlyak (2011), judicial requirement in foreclosure Mian, Sufi, and Trebbi (2015), categorization of sand states Choi, Hong, Kubik, and Thompson (2016), and state capital gain tax Gao, Sockin, and Xiong (2020).

4.2 Over-identification Test

So far, I have not been able to test whether the regression specification in Equation (9) is correct or not by an over-identification test (Hansen J test) primarily because I only have one instrument variable. In this subsection, I use the gravity model-based instrument ($\triangle_{99,05}\text{givNetExp}_m$) and its square as two separate instrumental variables for the growth rate of private-label mortgages in Equation (9) so that I can conduct an over-identification test. This test can provide suggestive evidence of whether the specification in Equation (9) is correct statistically.

Table (7) reports OLS, reduced-form, second-stage, and first-stage results in panels A, B, C, and D, respectively. The reduced-form estimates of two separate IVs show that two separate IVs are insignificant in columns (3) to (5). However, the two coefficients are large in magnitude (9.492 and -356.147) and have opposite signs. In columns (3) to (5) in panel D, the first-stage coefficient estimates are insignificant, and F-statistics are not large enough, both of which indicate weak IV concern. However, 2SLS coefficient estimates in Panel C are significantly positive and have magnitudes (0.437) almost identical to the 2SLS estimates (0.438) in Table (3) with only one IV. In addition, in column (1) of first-stage regression and column (6) of 2SLS regression, coefficients and standard errors of control variables in Table (6) with two IVs are almost identical to Table (6) with only one IV. The above results together suggest that the added IV of square of gravity model-based instrument ($\triangle_{99,05}\text{givNetExp}_m^2$) does not provide additional information because the new 2SLS coefficient estimates are almost identical to the original ones. Moreover, the square of gravity model-based instrument introduces a multicollinearity problem with the original IV of gravity model-based instrument ($\triangle_{99,05}\text{givNetExp}_m$), resulting in insignificant coefficients in the first-stage and reduced-form regressions. In other words, the square of gravity model-based instrument is redundant. I take advantage of this redundant IV to perform the over-identification test. Hansen J statistics are small enough, and p-values are large enough, especially in column (6) with full controls in Table (6) (0.403). These results suggest that my original specification with one instrumental variable in Equation (9) is correctly specified statistically.

5 Empirical Results: Against Speculation

In this section, I design five tests to show that credit expansion instead of speculation plays the dominant role in causing the house price cycle across metropolitan areas. In my first test, I demonstrate that net export growth does not cause government-sponsored enterprise mortgage growth during the boom period (1999-2005). This test addresses the speculation (demand) view because demand does not appear in this chapter mortgage type (GSEM). My second test validates the exclusion restriction of my IV strategy: net export growth causes house price growth in the boom period (99-05) only through the channel that it causes growth in private-label mortgages (non-jumbo). In particular, in a prior period (1991-1999) without substantial credit expansion, net export growth does not cause house price growth precisely because it does not cause growth in private-label mortgages (non-jumbo). Put differently, net export growth cannot directly cause house price growth via the demand channel without credit expansion in mortgages.

The following three tests are designed to address the potential concern from the "speculation" view that, even though credit expansion happens in the first place, speculation by borrowers may also dominate the mortgages by taking advantage of the credit expansion. This potential concern originated from the viewpoints of Kindleberger (1978) and Minsky (1986). Since government-sponsored enterprise mortgages (GSEMs) do not respond to net export growth (see Table (11)), they cannot explain the cross-metro housing cycle. I only distinguish speculation and credit expansion within the private-label mortgages (non-jumbo). I use the speculation measure by Gao, Sockin, and Xiong (2020), which is the "non-owner-occupied" home purchase mortgages. Following this logic, I use "owner-occupied" home purchase mortgages to measure pure credit expansion. First, I demonstrate that the cross-metropolitan variation in speculation can be largely explained by pure credit expansion. I use the residuals from this first specification as my measure of the growth rate of credit-independent speculation, which is part of the growth rate of the non-owner-occupied private-label mortgage (non-jumbo) that cannot be explained by the growth rate of owner-occupied private-label mortgage (non-jumbo). The second test shows that, compared to credit-independent speculation, credit expansion plays the dominant role in explaining the house price boom. Specifically, credit expansion explains twelve times the house price growth that speculation explains. The third test focuses on the prior period (91-99) where there is substantial differences in net export growth across metropolitan areas but no aggregate credit expansion. I show that, without credit expansion, speculation does not increase in response to local economic conditions caused by net export growth.

Before going into regression specification, it is helpful to see the general trend of credit expansion vs. speculation mortgage categories across time in Figure (12). First, in the boom period (99-05), the credit expansion (owner-occupied home purchase) is much higher in absolute dollar amount than

speculation (non-owner-occupied home purchase). Growth rates of both measures are stronger in the high net-export-growth areas. Such observations are consistent with my conclusion for tests 1 & 2 discussed above. For example, in the peak year of 2005, for the top quintile metropolitan areas based on net export growth, pure credit expansion is 67.0 Billion (07USD), whereas the speculation is only 13.9 Billion (07USD). In addition, in the prior period, there is not much difference in speculation (non-owner-occupied home purchase) between the high and low net-export-growth metropolitan areas. This observation is consistent with the conclusion in my third test above. Please note that for ease of calculation, all dependent and independent variables are annualized in this section.

5.1 Relevance: Private vs. Government Mortgages in Boom (99-05)

Motivated by the legal constraint that government-sponsored enterprise mortgages (GSEMs) cannot consider regional economic conditions in setting up mortgage rates (Hurst, Keys, Seru, and Vavra, 2016), I expect that GSEMs do not respond to net export growth and thus are not related to the differential housing boom and bust across metropolitan areas. However, due to peer effect from neighbors who use credit expansion in private-label mortgages (non-jumbo) (PLMNJ) for housing consumption, credit-qualified households may use government-sponsored enterprise mortgages (GSEMs) for their house consumption. In such cases, ignoring GSEMs can overestimate the impact of PLMNJ. This section addresses such concern by showing the irrelevance of GSEMs to the housing cycle. Here, I distinguished two effects of net export growth: realized effects vs. unrealized but predicted effects in the future. Realized effects are realized at the household level, including income growth, credit score increase, etc. These realized effects at the household level can be considered by GSEMs. I argue such realized effects are still small in the sample periods, so they cannot be picked up in the regression analysis. In a similar spirit, my argument against demand does not mean demand does not play any role. In fact, the realized effects above (caused by net export growth) can drive up housing demand in GSEMs, which seems not to show up in the regression. Overall, I only mean that credit expansion plays the dominant role. In contrast, unrealized but predicted effects in the future are forward-looking differences at the metropolitan level, including future average income growth, average credit score increase, average house price growth, etc. These unrealized but predicted effects associated with metropolitan areas can only be considered by PLMs, not GSEMs.

In addition, the irrelevance of GSEMs can help argue against the demand view of mortgage growth. Since government-sponsored enterprise mortgages (GSEMs) are still cheaper than corresponding non-jumbo private-label mortgages (PLMNJs) for credit-qualified households, mortgage demand driven by net export growth would show up in both GSEMs and PLMNJs. That is to say, we would see higher mortgage growth in high net-export-growth metros for both GSEMs and PLM-

NJs. However, in the empirical tests, we only see such a trend for PLMNJs rather than GSEMs. Therefore, such empirical evidence helps argue against the above speculation (demand) view.

Hypothesis 2 In the boom period 1999-2005, in cross-section, net export growth causes credit expansion in private-label mortgages (non-jumbo) (PLMNJ) rather than government-sponsored enterprise mortgages (GSEM).

Private Mortgage Growth in Boom Period (99-05) For PLMNJ, I estimate the following regression

$$\triangle_{99,05} Ln(PLMNJ_c) = \beta * \triangle_{99,05} \text{NetExp}_m + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c \tag{11}$$

where I use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for net export growth ($\triangle_{99,05}$NetExp$_m$).

Table (9) reports OLS, reduced-form, second stage, and the first stage of the results in panels A, B, C, and D. First, panel A column (1) shows the positive and significant impact of net export growth (99-05, annualized) on the growth of private-label mortgages (non-jumbo) (PLMNJ) (99-05, annualized), without any control. Through columns (2) to (4), the estimates show that the OLS coefficient of the next export growth is almost unaffected by the inclusion of the four categories of controls. Panel B shows that the reduced-form coefficients are significant and stable to the inclusion of various controls. Panel D show that the first-stage coefficients are quite stable across various specifications. In column (5) with all control variables, the first-stage clustered Kleibergen-Paap F-statistic is 17.11, and the Montiel Olea-Pflueger Efficient F-Statistic is 16.83. Thus, it is very unlikely that my estimates are biased by a weak instrument. Like reduced-form estimates, the 2SLS estimates in Panel C are statistically significant at a one percent level and quite stable across various specifications. Besides my two F-Statistics, I also confirm that 2SLS is close to OLS estimates throughout various specifications, where my baseline results in column (5) indicate a ratio of $11.715/6.786 = 1.726$. Table (8) Column (1) reports the first-stage results where most controls are not significantly correlated with net export growth. This result indicates that net export growth is orthogonal to most local initial economic conditions.

Let us focus on the economic meaning in 2SLS with controls in Table (8). According to column (5) with all controls, one standard deviation in cross-sectional difference in annualized net export growth results in $0.20\% \times 10.99 = 2.20\%$ difference in annualized private-label mortgages (non-jumbo) (PLMNJ) growth, translating into 13.22% difference at the county level from 1999 to 2005. One standard deviation in cross-section difference in annualized PLMNJ growth is 8.14%, translating into 48.82% from 1999 to 2005. A quick calculation means that, from 1999 to 2005, one standard deviation in net export growth can explain $13.22\%/48.82\% = 27.07\%$ of one standard deviation in PLMNJ growth.

Likewise, one standard deviation in cross-section housing supply elasticity (HSE) can explain $1.217 * 0.010 = 1.217\%$ difference in annualized PLMNJ growth, translating into 7.30% difference

at the county level from 1999 to 2005. A quick calculation implies that, from 1999 to 2005, one standard deviation in HSE can explain 7.30%/48.82% = 14.95% of one standard deviation in PLMNJ growth. In comparison, the explanatory power of my net export growth (27.07%) is around twice that of housing supply elasticity (14.95%), a major factor documented in the literature (Mian and Sufi, 2011; Saiz, 2010).

Private vs. Government Mortgage Growth in Boom Period (99-05) Recall that I predict that GSEMs will not expand more in the boom period because GSEMs do not consider differences in regional economic growth, including the net export growth.

Table (11) reports OLS, reduced-form, second stage, and the first stage of the stacked regression of GSEM and PLMNJ in panels A, B, C, and D. Regression specifications are similar to equation (11) except that (1) the dependent variable stacks growth of both GSEM and PLMNJ; and (2) independent variables interact with both dummy variables for GSEM and PLMNJ. First, panel A shows that the OLS coefficients of net export growth (99-05, annualized) interacted with the PLMNJ dummy are precisely the same between stacked regressions and previous unstacked ones. It is worth noting that OLS coefficients of net export growth on GSEM growth are significant. In sharp contrast, their reduced-form counterparts are statistically insignificant across various specifications in panel B. This contrast highlights the necessity of instrumental variables since OLS estimates might pick up some non-exogenous part of net export growth. For example, due to the long-term trend of supply-side technology advances unique to US industries beyond eight other high-income countries, the booming local industries expand both exports and employment. Since such technological advances result in employment growth, demand for GSEMs grows. In sharp contrast, my testing boom period is relatively short, and exogenous change in net export may not reflect the above long-term trend of supply-side advances. Related companies are unlikely to expand their employment substantially. Due to legal constraints (Hurst, Keys, Seru, and Vavra, 2016), GSEMs cannot consider small differences across geographic areas caused by exogenous net export growth. My reduced-form estimates are consistent with GSEMs ignoring such differences.

Consistent with the above, the 2SLS estimates of the interaction term between net export growth and GSEM growth are mostly insignificant, especially when more controls are added. Let us turn our attention to the coefficient equality test of two interaction terms (net export growth interacted with Dum_PLMNJ and Dum_GSEM) in 2SLS in Table (10). For all specifications, the chi-square statistics are large, and p-values are below 0.05, meaning the two coefficients are statistically different. In addition, the insignificant and close-to-zero coefficient of interaction term housing supply elasticity $\times$ Dum_GSEM confirms that GSEM do not consider differences in land availability for housing across counties in metropolitan areas.

5.2 Exclusion Restriction: Prior (91-99) vs. Boom (99-05)

This subsection provides evidence supporting the exclusion restriction of our IV approach: net export growth impacts house prices only via its impact on PLMNJs. This subsection is designed to address the concern that net export growth directly causes stronger house price growth in the high net-export-growth areas, or indirectly contributes to house price increases through a different channel. In the above two situations, the IV strategy can overestimate the impact of PLMNJ growth.

My test focuses on the prior period (1991-1999) when there is no credit expansion and the funding cost of private-label mortgages (non-jumbo) (PLMNJs) is still high. In the prior period, two factors prevent PLMNJs from expanding more in the high net-export-growth areas: (1) government-sponsored enterprise mortgages (GSEMs) dominate the mortgage markets because of low rates due to implicit government insurance (Sherlund, 2008), and (2) GSEMs do not consider differences in regional economic growth (Hurst, Keys, Seru, and Vavra, 2016). Therefore, the prior period can provide evidence supporting the exclusion restriction of IV: net export growth impacts house prices only via its impact on PLMNJs. In other words, I can show that net export growth did not cause higher growth in PLMNJs in the high net-export-growth areas in the prior period (1991-1999). Consequently, it did not increase house prices in these areas. I choose a long period from 1991 to 1999 since I want to include two events: (1) the North American Free Trade Agreement enacted in 1994 and (2) the World Trade Organization established in 1995. These two events make the prior period an ideal setting for a placebo test: there is a substantial divergence in net export growth across metropolitan areas but no credit expansion at the aggregate level. Again, here I distinguished realized effects vs. unrealized but predicted effects in the future. My argument against the demand (speculation) view does not mean demand does not play any role here. In fact, the realized effects on household income and credit scores (caused by net export growth) can drive up housing demand in PLMNJs in the prior period. However, this demand channel does not show up in the empirical tests. Overall, I only mean that net export growth does not directly cause house price growth significantly.

Hypothesis 3 In cross-section, net export growth causes house price growth in the boom period (99-05) because it causes growth in private-label mortgage (non-jumbo) (PLMNJ). Conversely, net export growth does not cause house price growth in the prior period (91-99) because it does not cause growth in private-label mortgage (non-jumbo) (PLMNJ).

In the first step, I test the relationship between net export growth and house price growth in both periods. Table (13) reports OLS, reduced-form, second stage, and first stage of the stacked regression of house price growth in the prior period (91-99) and boom period (99-05) in panel A, B, C, and D, respectively. Regression specifications are similar to equation (11) except that (1) the dependent variable stacks house price growth in the prior period (91-99) and the boom period

(99-05) and (2) independent variables interact with dummy variables for two periods.[46][47] First, panel A shows that the OLS coefficients of net export growth (99-05, annualized) are positive and statistically significant at a one percent level. However, the OLS estimates for net export growth (91-99, annualized) are insignificant. The same trend applies to reduced-form estimates in panel B and 2SLS estimates in panel C. First-stage estimates in panel D show the strong positive correlation between the net export growth and its gravity model-based instrumental variable is quite stable for both periods. In column (5) with all control variables, the first-stage clustered Kleibergen-Paap F-statistic is 29.87. And the Montiel Olea-Pflueger Efficient F-Statistic is 29.76. Thus, my estimates are very unlikely biased by a weak instrumental variable. Let us turn our attention to the coefficient equality test of net export growth in prior and boom periods in 2SLS in Table (12). In prior period (91-99), as more controls are added, coefficients for net export growth are insignificant in column (3)-(5). The chi-square statistics are large, and p-values are below 0.05 for all specifications, meaning the two coefficients are statistically unequal. Taken together, the results from the above two tables verify that house price growth between 1991 and 1999 is not affected by net export growth.

As a second step, I test the relationship between net export growth and growth in private-label mortgages (non-jumbo) (PLMNJ) in both periods. Table (15) reports OLS, reduced-form, second stage, and first stage of the stacked regression of PLMNJ in the prior period (91-99) and boom period (99-05) in panel A, B, C, and D, respectively. [48][49] First, panel A shows that the OLS coefficients of net export growth (99-05, annualized) are positive and statistically significant at a one percent level and are stable across specifications. However, the OLS estimates for net export growth (91-99, annualized) are generally insignificant. The same trend applies to reduced-form estimates in panel B and 2SLS estimates in panel C. First-stage estimates in panel D show that the strong positive correlation between the net export growth and its gravity model-based instrumental variable is quite stable for both periods. In column (5) with all control variables, the first-stage clustered Kleibergen-Paap F-statistic is 34.08, and the Montiel Olea-Pflueger Efficient F-Statistic

[46] Please note that I require counties to be consistently covered by the HMDA database since 1991. Because of the smaller inclusion of metropolitan areas and the availability of the housing price index in the early years, my sample size is reduced to $672 \times 2 = 1344$ counties in columns (1)-(2) and further reduced to $599 \times 2 = 1198$ in column (5) due to data availability of housing supply elasticity.

[47] In the prior period, I winsorize the net export growth and its GIV at the 3% and 97%, since a few outliers (based on GIV net export growth) can make the coefficient of net export growth negative and significant. Even though the negative and significant coefficient does not violate my conclusion, I winterize the data to show the results for most observations. In the Appendix Section A.3, I will show more details.

[48] Please note that I require counties to be consistently covered by the HMDA database since 1991. Because of the smaller inclusion of metropolitan areas in the early years, my sample size is reduced to $702 \times 2 = 1404$ counties in columns (1)-(2) and further reduced to $624 \times 2 = 1248$ counties in column (3)-(5) due to data availability of housing supply elasticity.

[49] In the prior period, I drop three outliers, since these three outliers can make the coefficient of net export growth negative and significant. Even though the negative and significant coefficient does not violate my conclusion, I drop three outliers to to show the general results for most observations. In the Appendix Section A.3, I will show more details.

is 34.74. Thus, my estimates are very unlikely biased by a weak instrumental variable. Let us turn our attention to the coefficient equality test of net export growth in prior and boom periods in 2SLS in Table (14). For specifications in columns (1)-(5), the chi-square statistics are large, and p-values are below 0.05, meaning the two coefficients are statistically different. Taken together, results from the above two tables verify that PLMNJ, between 1991 and 1999, is not affected by net export growth. It is also interesting that the specification in column (5) shows that PLMNJ growth between 1991 and 1999 is unaffected by housing supply elasticity after considering industry controls.

In summary, my empirical evidence in this subsection confirms hypothesis 3, verifying the exclusion restriction in my IV strategy. That is to say, net export growth can impact house prices only through its impact on private-label mortgages (non-jumbo).

5.3 Credit Expansion Causes Speculation

In this subsection, I show that cross-metropolitan variation in speculation is caused by credit expansion. I use the following 2SLS specification:

$$\triangle_{99,05} Ln(PLMNJ_NonOwn_c) = \underbrace{\beta * \triangle_{99,05} Ln(PLMNJ_Own_c)}_{\text{Credit-Induced Speculation}} + \underbrace{\gamma * \boldsymbol{Controls_c} + \epsilon_c}_{\text{Credit-Independent Speculation}} \quad (12)$$

The left-hand-side dependent variable $\triangle_{99,05} Ln(PLMNJ_NonOwn_c)$ is the growth rate of the dollar amount of non-owner-occupied private-label mortgages (non-jumbo) at county c 99-05 and the key independent variable $\triangle_{99,05} Ln(PLMNJ_Own_c)$ is the growth rate of the dollar amount of owner-occupied private-label mortgages (non-jumbo) at county c 99-05. I use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05} Ln(PLMNJ_Own_c)$. Controls, weights, and standard errors are the same as Eq (11).

Table (16) reports OLS, reduced-form, second stage, and the first stage of the results. First, panel A shows the positive and significant impact of owner-occupied private-label mortgages (non-jumbo) (PLMNJ) (99-05, An) as credit expansion on the growth of non-owner-occupied PLMNJ as speculation. This coefficient is quite stable to the inclusion of various controls. Panel B reports the reduced-form estimates and shows that gravity-model-based net export growth increases the speculation. First-stage estimates in panel D show that the strong positive correlation between the GIV and net export growth is quite stable across various specifications. In column (5) with all control variables, the first-stage clustered Kleibergen-Paap F-statistic is 10.34, and the Montiel Olea-Pflueger Efficient F-statistics is 10.28. Thus, it is very unlikely that my estimates are biased by weak instruments. The 2SLS estimates in panel C are statistically significant at a one percent level and quite stable across various specifications. Besides my two F-statistics, I also confirm that 2SLS is close to OLS estimates throughout various specifications, where my baseline results in column (5) indicate a ratio of $1.190/1.119 = 1.063$, much lower than nine. In terms of economic

meaning in 2SLS with full controls, one standard deviation of 6-year PLMNJ (owner-occupied, credit expansion) growth can cause $1.19 * 0.46 = 55.18\%$ increase in 6-year PLMNJ (non-owner-occupied, speculation) growth, which is $55.18\%/107.49\% = 51.34\%$ of one standard deviation of PLMNJ (non-owner-occupied) growth. Therefore, I can conclude that cross-metropolitan variation in speculation is largely driven by exogenous variation in credit expansion.

5.4 Credit-Independent Speculation vs Pure Credit Expansion

In this subsection, I use the decomposition from the first test to get credit-independent speculation. Then I show that, compared to the dominant role of pure credit expansion measured by owner-occupied private-label mortgages (non-jumbo), credit-independent speculation can only explain a tiny portion of housing price boom.

To be specific, I use the following regression specification:

$$\triangle_{99,05} Ln(HPI_c) = \beta * \triangle_{99,05} Ln(PLMNJ_Own_c) + \gamma * \text{Credit-Independent Speculation} + \boldsymbol{Controls_c} + \epsilon_c \quad (13)$$

One key independent variable $\triangle_{99,05} Ln(PLMNJ_Own_c)$ is the growth rate of the dollar amount (07USD) of owner-occupied private-label mortgage (non-jumbo) (PLMNJ) as pure credit expansion at county c 99-05. This variable is instrumented by the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$). The other key independent variable Credit-Independent Speculation is derived from the regression Eq (12), which is the part of growth rate of the non-owner-occupied private-label mortgage (non-jumbo) that cannot be explained by the pure credit expansion.

Table (17) reports OLS, reduced-form, second stage, and the first stage of the results. First, panel A shows the positive and significant impact of owner-occupied private-label mortgages (non-jumbo) (PLMNJ) (99-05, An) as pure credit expansion on the house price growth. Panel B reports the reduced-form estimates and shows that GIV net export growth explains the housing price growth. First-stage estimates in panel D show that the strong positive correlation between the GIV net export growth and the growth rate of pure credit expansion. In contrast, credit-independent speculation cannot explain such growth. In column (5) with all control variables, the first-stage clustered Kleibergen-Paap F-statistic is 10.36, and the Montiel Olea-Pflueger Efficient F-statistics is 10.30. Thus, it is very unlikely that my estimates are biased by weak instruments. The 2SLS estimates in panel C are statistically significant and quite stable across various specifications. In term of the economic meaning in 2SLS with full controls, one standard deviation of six-year owner-occupied PLMNJ growth can cause $0.454 * 0.463 = 21.05\%$ house price growth. In comparison, one standard deviation of six-year credit-independent non-owner-occupied PLMNJ growth can cause $0.018 * 0.947 = 1.70\%$ house price growth. That is to say, the pure credit expansion can explain more than twelve times ($21.05\%/1.70\% = 12.35$) of the house price growth that is explained by credit-independent speculation.

Results together show that home consumption demand ("owner-occupied") induced by credit expansion plays a much more important role than speculation ("non-owner-occupied"). Therefore, I conclude that credit expansion can explain the majority of the house price growth, while credit-independent speculation can only explain a tiny part of the house price growth. This conclusion is slightly different from the view by Kindleberger (1978); Minsky (1986).[50]

5.5 Speculation Cannot Grow Without Credit Expansion

In this subsection, I show that speculation cannot grow without credit expansion. Recall that in Eq (12), I have shown that net export growth causes the credit expansion that largely explains the cross-sectional variation of the speculation. To further show that credit expansion is a necessary condition for speculation, I focus on the prior period (91-99) when there is no aggregate credit expansion and I show that even the tradable sector growth cannot cause credit expansion.

Specifically, I use the following regression specification:

$$\begin{aligned}\triangle_{91,99}\&\triangle_{99,05}Ln(PLMNJ_NonOwn_c) = \beta_{91,99} * \triangle_{91,99}\text{NetExp}_m \times Dum_{91,99} + \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} \\ + \gamma_{91,99} * \boldsymbol{Controls_c} \times Dum_{91,99} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \epsilon_{period,c}\end{aligned} \tag{14}$$

The left-hand-side dependent variable $\triangle_{91,99}\&\triangle_{99,05}Ln(PLMNJ_NonOwn_c)$ is the stacked growth rate of the dollar amount (07USD) of non-owner-occupied private-label mortgages (non-jumbo) (PLMNJ) at county c 91-99 and 99-05, respectively. Again, the key independent variable, net export growth, is instrumented by the gravity model-based IV in two periods, respectively.

Table (18) reports OLS, reduced-form, second stage, and the first stage of the results. First, OLS regressions in panel A show that net export growth is positively correlated with the speculation in the boom period (99-05) but not in the prior period (91-99). The same pattern shows up in the reduced-form regressions. First-stage estimates in panel D and E show the strong positive correlation between the net export growth and its GIV in both periods. In column (5) with all control variables, the first-stage clustered Kleibergen-Paap F-statistic is 73.77 in the prior (91-99) and 17.54 in the boom (99-05), and the Montiel Olea-Pflueger Efficient F-statistics is 77.33 and 17.07. The 2SLS estimates in panel C show that net export growth causes speculation in the boom period but bears no relation in the prior period (91-99). The above empirical results are more consistent with the view that net export growth can cause speculation only if there is aggregate credit expansion. Put together, these three tests are more consistent with the view that credit expansion is a necessary condition for speculation.

At this stage, I summarize the empirical evidence presented thus far that argues against the "speculation" (or demand) view.

First, in Section 5.1, I use government-sponsored enterprise mortgages (GSEMs) as a control

[50]This conclusion only refers to the first mortgage market. Though out of the scope of this paper, it might be the case that speculation by investors plays an important role in the secondary mortgage market.

group to argue that the differentially higher growth rate in private-label mortgages (PLMs) in the high net-export-growth areas is not driven by demand. Since the mortgage rate of GSEM is always lower than PLM for the same borrower (Justiniano, Primiceri, and Tambalotti, 2022; Sherlund, 2008), the demand-driven mortgage by borrowers shall show up in the differentially higher growth in GSEMs in the high net-export growth areas. However, this demand-drive conjecture is not supported by empirical evidence. In fact, GSEMs experience no difference in growth rate between the high and low net-export-growth areas. Critical readers might think borrowers cannot know the differences between GSEM and PLM and only the lenders determine the choice of GSEM or PLM. Since government-sponsored enterprises (Fannie Mae and Freddie Mac) charge insurance fees and provide credit risk isolation, borrowers are more willing to originate GSEM than PLM for the same borrower whenever possible.

Second, in Section 5.2, I use the prior period (1991-1999) to show that differentially higher growth in PLMs in the high net-export-growth areas is not likely driven by demand in response to net export growth. In the prior period (1991-1999), where we know there is no aggregate credit supply, the high net-export-growth areas did not experience a higher growth rate in PLMs or GSEMs. This prior period fact shows at least one scenario that, without credit expansion, net export growth does not drive demand for mortgages (GSEMs and PLMs).

Third, Section 5 is designed to address the potential concern from the "speculation" view that, even though credit expansion happens in the first place, speculation by borrowers may also play an important role by taking advantage of the credit expansion. Following Gao, Sockin, and Xiong (2020), I measure speculation by the "non-owner-occupied" private-label mortgages for home purchase. Following the same logic, I measure credit expansion by "owner-occupied" private-label mortgages for home purchases. My first test shows that the cross-metropolitan variation in speculation can be mostly explained by credit expansion. The second test shows that, compared to credit-independent speculation, credit expansion plays the dominant role in explaining the house price boom. Third, in the prior period (1991-1999) where there were substantial differences in net export growth across metropolitan areas but no credit expansion at the aggregate level, I show that speculation does not respond to the net export across metropolitan areas.

6 Model

In this section, I build a simple theoretical model to illustrate how credit expansion alone can reconcile the two seemingly opposing empirical facts: the correlation between income growth and mortgage growth is negative across ZIP codes within metropolitan areas but positive across metropolitan areas. Thus, this section addresses the potential concern that two seemingly opposing empirical facts may result in internal conflicts within the "credit expansion" view. I can summarize the

model intuition as follows. First, when the funding cost of private-label mortgages declines within metropolitan areas, lenders can extend cheap credit to low-income borrowers traditionally excluded from the mortgage markets due to usury law. Since households with sharp rising wealth tend to self-relocate to rich ZIP codes and households with sharp declining wealth tend to self-relocate to poor ZIP codes, income growth in poor ZIP codes is usually lower than in rich ZIP codes, even income growth distribution is the same across all occupations, industries and initial wealth level.[51] Thus, the negative correlation between income growth and mortgage growth within metropolitan areas means that poor ZIP codes see higher mortgage growth. Second, given the above within-metropolitan fact, poor ZIP codes in the high net-export-growth areas experience relatively higher private-label mortgage growth than those in the low net-export-growth areas due to the differences in net export growth. Thus, the correlation between income growth and mortgage growth is positive across metropolitan areas. In summary, credit expansion induced by net export growth alone can reconcile empirical facts within and across metropolitan areas.

6.1 Mortgage First & Secondary Market Structure

First, let us describe the market structure in the mortgage first and secondary markets in the United States. In the mortgage first market, each household enters into a mortgage contract with a lender separately if the mortgage application is approved. Mortgage lenders include deposit institutions (commercial banks, thrifts, and credit unions) and non-deposit institutions (independent mortgage companies).

After origination, most mortgages in the United States are sold to a secondary market rather than staying on lenders' balance sheets. For example, about 80 percent of mortgages were sold and securitized between 2004 to 2006 (Keys, Piskorski, Seru, and Vig, 2013). In the secondary mortgage market, when the mortgages are qualified for purchase by government-sponsored enterprises (GSEs include Fannie Mae, Freddie Mac, and Ginnie Mae), lenders sell these mortgages to GSEs. These mortgages are called government-sponsored enterprise mortgages (GSEM). GSEs insure GSEMs against loss of principal and interest via charging insurance fees. GSEs then package these mortgages into mortgage-backed securities (MBS), collateralized mortgage obligations (CMO), and other types of securities. Eventually, these securities are sold to investors with relatively low-risk tolerance. Typical investors include mutual funds, pension funds, and insurance companies. Because of lower funding costs due to implicit guarantees by the government and the large scale of the GSEM market (Hurst, Keys, Seru, and Vavra, 2016), lenders will first choose to sell mortgages to GSEs if these mortgages are qualified. It is important to note that GSEs do not consider regional economic conditions in setting mortgage rates.

However, if mortgages are not qualified as GSEMs, lenders will sell mortgages to private is-

[51] This within-metropolitan cross-ZIP codes self reallocation is in the spirit of "voting by foot" by Tiebout (1956)

suers. Typical private issuers include large commercial banks, large mortgage companies, investment banks, and real estate investment trusts (REITs). These mortgages are called private-label mortgages (PLMs). Contrary to GSEs, private issuers consider regional economic conditions when setting mortgage rates and do not provide insurance against principals or interests. Like GSEs, private issuers package these mortgages into MBS, CMO, and other types of securities (usually by creating special purpose vehicles (SPV)). Ultimately, these securities are sold to investors with relatively high risk tolerance who bear the default risk of the underlying mortgages. Typical investors include hedge funds, investment banks, and REITs.

Now, let us describe each player's required returns and fees in the mortgage first and secondary market. Figure 3 depicts the entire market. Since the lenders now play the role of originate-to-distribute, they only charge transaction fees (underwriting services, etc.) TF_L (as a percentage of house value). In the GSEM market, since GSEs provide insurance and bear the default risk, they charge an insurance fee of GIR (as a percentage of house value). They also charge a transaction fee for securitization TF_I (as a percentage of house value). Low-risk investors only charge a (close to) risk-free rate r_f since they are fully insured against any potential loss. I define the funding cost of GSEM as $FCR_G \coloneqq r_f + GIR$. In sum, to fund a mortgage in the GSEM market, the total required return is $1 + FCR_G + TF$, where $TF \coloneqq TF_L + TF_I$.

In the PLM market, however, private issuers only charge transaction fees for securitization TF_I without providing insurance. As default risk is eventually borne by high-risk investors, they charge risk-adjusted return (defined as funding cost for PLM market) $FCR_P \coloneqq r_f + \delta(g_u - g_d)$. δ is the risk-averse parameter, g_u and g_d are household income growth rates for up-state and down-state, respectively. Thus $g_u - g_d$ approximates the uncertainty in the household income growth, and $\delta(g_u - g_d)$ approximates the risk premium in mortgage rate. In sum, to fund a mortgage in the PLM market, the total required return is $1 + FCR_P^* + TF$, where again $TF \coloneqq TF_L + TF_I$.

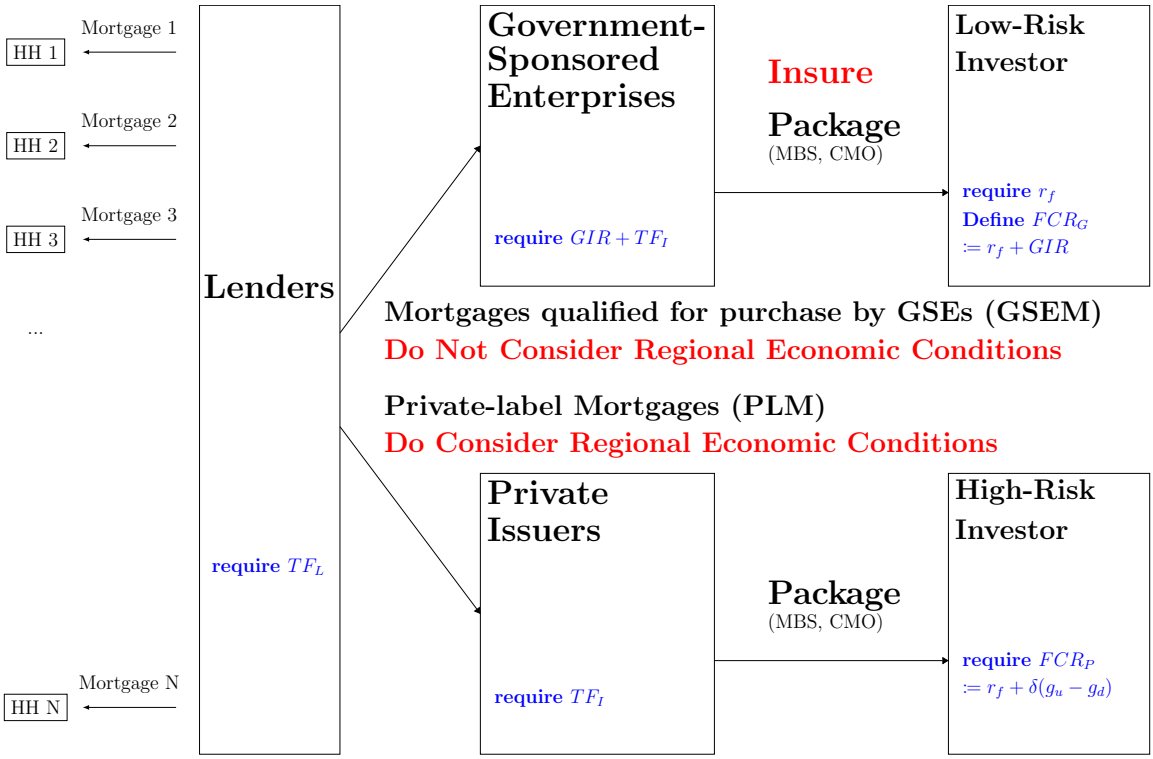


Figure 3: Mortgage First and Secondary Markets

6.2 Basic Setup

Now let us set up the environment of the model. There are two areas in the economy: an area with high net export growth (HNEG area) and an area with low economic growth (LNEG area). In the model, I assume net export growth can increase local employment growth, household income growth, and population growth (including migration).[52] Thus, the HNEG area has an additional positive employment growth rate ($+\Delta e$), an additional positive household income growth rate ($+\Delta g$), and an additional positive population growth rate ($+\Delta p$) while the LNEG area has an additional negative employment growth rate ($-\Delta e$), an additional negative household income growth rate ($-\Delta g$), and an additional negative population growth rate ($-\Delta p$). Each area has both secondary mortgage markets described above, GSEM and PLM markets.

In each area, there are an infinite number of homogeneous lenders who play the role of originate-to-distribute. The market players and their required returns and fees are described in the previous section. In each area, there are an infinite number of heterogeneous households that differ only in initial income-to-house ratio $i_0 := \frac{I_0}{H_0}$, where I_0 and H_0 are the household income and house value in time 0. To make households from both areas comparable, I assume the household income-to-house

[52] For the economic literature of why net export can increase these local economic conditions and my verification tests, see Section A.5

ratios in time 0 follow the same distribution in both the LNEG area and the HNEG area.[53] In addition, I assume the house price is the same within both the HNEG area ($H_{H,0}$) and the LNEG area ($H_{L,0}$) because my focus is the cross-area comparison of the market share of GSEM and PLM. This assumption effectively avoids the unnecessary complexity of income-house matching within each area. With the simplified assumptions above, I can achieve a closed-form solution that has intuitive economic interpretation and graph demonstration.

There are two periods ($t = 0$ & 1) in the model. In the first period ($t = 0$), each household starts with no initial wealth and earns income I_0, which is observable to the lenders. For the simplicity of modeling, I assume the household consumes all income I_0 in time 0. In the first period ($t = 0$), each household enters into a mortgage contract with a lender by pledging the house as collateral, if their income-to-house ratio i_0 is qualified for the mortgage. I assume utility-maximizing households get higher utility from living in a house than renting an apartment. Thus households will get mortgages as long as they can. Otherwise, they rent an apartment and pay rent at the end of time 1.

In the second period ($t = 1$), there is uncertainty at the individual household level: the income growth could have two states with equal probability ($\frac{1}{2}$)[54]: an up-state with income growth rate g_u and a down-state with income growth rate g_d, where $g_u > g_d$. In the up-state, the household income growth rate is $g_u + \Delta g$ in the HNEG area and $g_u - \Delta g$ in the LNEG area. Similarly, in the down-state, the household income growth rate is $g_d + \Delta g$ in HNEG area and $g_d - \Delta g$ in LNEG area. Since uncertainty exists only at the individual household level, house prices only vary with area-specific economic growth rate, meaning $H_0(1 + \Delta g + \Delta e + \Delta p)$ in HNEG area and $H_0(1 - \Delta g - \Delta e - \Delta p)$ in LNEG area. I further assume that in time 1, the household income is sufficient to pay for the mortgage amount in the up-state, while not sufficient in the down-state. This two-state assumption is meant to be simple so that I can get a closed-form solution of mortgage rate with easy economic interpretation. With this assumption, to make the mortgage rate linearly decrease with income-to-house ratio i_0 (which is realistic in practice), I further assume that households would forgo both the house (as collateral) and disposable income (as a penalty) to the lender in case of default. The loss of disposable income in the down-state is motivated by three reasons. First, the foreclosure sale of the house usually cannot cover the entire loss of the mortgage, especially during the housing crisis. Second, mortgages in thirty-nine U.S. states are recourse loans, meaning lenders may be able to collect on a debt of households by obtaining a deficiency judgment when proceedings from a foreclosure sale are not enough to cover the loss of a mortgage (see Ghent and Kudlyak (2011)). Third, the standard mortgages in the United States have a time to maturity of 30 years such that some portion of payments is paid before default.

[53] It is important to note that I do not assume the initial house prices are the same in the HNEG area and the LNEG area.

[54] Equal probability of two states are meant to be simple for illustration purpose. For a general probability $p \in [\frac{1}{2}, 1]$ for up-state, my conclusions still hold. See Appendix A.6.2 for proof.

Next, I will focus on the break-even conditions in the following three situations: (1) GSEM in both HNEG area and LNEG area, (2) PLM in HNEG area, and (3) PLM in LNEG area. Since lenders only play the role of originate-to-distribute, the break-even condition includes required returns and fees by all parties in the market depicted in Figure (3).

6.3 GSEM in Both HNEG area and LNEG area

According to the analysis by Hurst, Keys, Seru, and Vavra (2016), the mortgage evaluation and pricing system by GSEs does not consider differences in regional economic conditions. In other words, the following two decisions are made without regional economic conditions: (1) whether a mortgage can be bought by GSE as an agency mortgage and (2) what the mortgage rate is set on the loan. I model these features as that GSEs ignore differences in employment growth rate (Δe), household income growth rate (Δg), and population growth rate (Δp) in both HNEG area and LNEG area. In addition, GSEs require an insurance premium GIR in equilibrium so that GSEs charge $GIR + R_G$ in the non-default case while insuring against any loss in the default case.

Due to legal constraints that prevent considering differences in regional economic conditions, lenders could only perceive the household disposable income in time 1 would be $I_0(1 + g_u - c)$ in up-state with probability $\frac{1}{2}$ and $I_0(1 + g_d - c)$ in down-state with probability $\frac{1}{2}$.[55] The full mortgage amount is H_0 and mortgage payment at time 1 is $H_0(1 + FCR_G + R_G + TF)$. FCR_G is the funding cost rate, which includes the insurance rate GIR required by GSEs and the risk-free rate r_f required by low-risk investors. R_G is the return paid to GSEs in up-state (when there is no default) to compensate for the potential loss in down-state. TF is the sum of transaction fees required by lenders (TF_L) for mortgage processing and by GSEs (TF_I) for packaging mortgages.

To make the model simple but still contains uncertainty in mortgage decision-making, I assume (1) all households can pay the full mortgage amount in up-state and (2) all households default in down-state. The first assumption means that even a household with the lowest income can pay the full mortgage amount in up-state in time 1: $\underline{I_0}(1 + g_u - c) \geq H_0(1 + FCR_G + R_G + TF)$. This inequality implies

$$\frac{\underline{I_0}}{H_0} \geq \frac{1 + FCR_G + R_G + TF}{1 + g_u - c} := \underline{i_G}$$

The second assumption means that even a household with the highest income could default in down-state in time 1: $\overline{I_0}(1 + g_d - c) < H_0(1 + FCR_G + R_G + TF)$. This inequality implies

$$\frac{\overline{I_0}}{H_0} < \frac{1 + FCR_G + R_G + TF}{1 + g_d - c} := \overline{i_G}$$

Therefore, I only assume that households have income-to-house ratio in time 0: $i_0 \in [\underline{i_G}, \overline{i_G})$. This assumption matches the practice. For the households that would default with probability 1 ($i_0 < \underline{i_G}$

[55] In this section, for simplicity, I omit the subscript of HNEG area and LNEG area in I_0 and H_0 since ultimately only i_0 matters and i_0 has same distribution among households in both areas.

in my setting), lenders would not grant mortgages at all. For the households that would not default at all ($i_0 > \overline{i_G}$ in my setting), they could always get mortgages by paying the lowest rate (insurance rate GIR plus risk-free rate r_f), thus having no variation in mortgage rates. In either case, there is no economic trade-off in mortgage lending that is key to my model demonstration. I purposefully ignore these two extreme groups of households to make my model simple and focus my attention on the key economic mechanism of interest.

Now let us focus on the key break-even condition for the GSEM market in both HNEG area and LNEG area: the expected payoff from households shall be enough to compensate the required return of ultimate low-risk investors ($r_f * H_0$), the insurance required by GSEs ($GIR * H_0$), and the transaction fees ($TF * H_0$):

$$\tfrac{1}{2}H_0 \underbrace{(1 + FCR_G + R_G + TF)}_{MR^*_G} + \tfrac{1}{2}[\underbrace{RR * H_0}_{\text{Recovery Value}} + \underbrace{I_0(1 + g_d - c)}_{\text{Disposable Income}}] = H_0(1 + FCR_G + TF) \quad (15)$$

The first term on the left-hand-side of the inequality is the product of probability in up-state ($\frac{1}{2}$) and the payment of the mortgage full amount $H_0(1 + FCR_G + R_G + TF)$. The second term is the product of probability in down-state ($\frac{1}{2}$) and the sum of the following two terms: (1) the recovery value of the mortgage from foreclosure sale from foreclosure sale $RR * H_0$ and (2) the disposable income $I_0(1 + g_d - c)$. A perfectly competitive market means that the mortgage rate MR^*_G is set to make the equality hold for each i_0 ex-ante :

$$\begin{aligned} MR^*_G &\coloneqq 1 + FCR_G + R_G + TF \\ &= [2(1 + FCR_G + TF) - RR] - (1 + g_d - c) * i_0 \\ &= C^*_G - K^*_G * i_0 \end{aligned} \quad (16)$$

The above mortgage rate has very clear economic interpretations. The mortgage rate in GSEM (in a non-default state) is positively related to funding cost ($FCR_G = r_f + GIR$, where r_f is the risk-free rate required by low-risk investors and GIR is the insurance rate required by GSEs in equilibrium) and the transaction fee required by lenders for mortgage processing and GSEs for packaging (securitization). In addition, the mortgage rate is negatively related to the recovery value of the mortgage from foreclosure sale (RR) and the disposable income in the default case ($(1 + g_d - c) * i_0$), since these two proceeds are used to cover loss given default.

6.4 PLM in HNEG area

In the GSEM market, since GSEs provide insurance against principal and interest, low-risk investors only require a risk-free rate as compensation. However, in the PLM market, lenders would consider regional economic conditions and high-risk investors would require a higher return to compensate for default risk due to a lack of insurance. I model these features in the following ways. First, PLM market (lenders, private issuers, and high-risk investors) would consider differences in regional

employment growth rate, household income growth rate, and population growth rate in both up-state and down-state. Second, high-risk investors would require a higher return $FCR_P^* = r_f + \delta * (g_u - g_d)$, where δ is a risk-averse parameter and $(g_u - g_d)$ (difference in household income growth rate for two states) is a measure of risk.

By observing household income $I_{H,0}$ at time 0, lenders know the household disposable income in time 1 would be $I_{H,0}(1 + g_u + \Delta g - c)$ in up-state with probability $\frac{1}{2}$ and $I_{H,0}(1 + g_d + \Delta g - c)$ in down-state with probability $\frac{1}{2}$. The full mortgage amount is $H_{H,0}(1 + FCR_P^* + R_{P,H} + TF)$. FCR_P^* is the funding cost required by high-risk investors in equilibrium because of default risk due to a lack of insurance. $R_{P,H}$ is the premium that goes to high-risk investors in up-state to compensate for the potential loss in down-state. TF is the transaction fee rate required by lenders for mortgage processing and by private issuers for packaging (securitization).

To make the model simple but still contains uncertainty in mortgage decision-making, I assume (1) all households can pay the full mortgage amount in up-state and (2) all households default in down-state. I get the following two boundary conditions for income-to-house ratios[56] :

$$\frac{\underline{I_{H,0}}}{H_{H,0}} \geq \frac{1 + FCR_P^* + R_{P,H} + TF}{1 + g_u + \Delta g - c} := \underline{i_{P,H}}$$

$$\frac{\overline{I_{H,0}}}{H_{H,0}} < \frac{1 + FCR_P^* + R_{P,H} + TF}{1 + g_d + \Delta g - c} := \overline{i_{P,H}}$$

Again, I only assume that households have income-to-house ratio in time 0: $i_0 \in [\underline{i_{P,H}}, \overline{i_{P,H}})$.

Now let us focus on the key break-even condition for the PLM market in HNEG area: the expected payoff from households shall be enough to compensate the required return of ultimate high-risk investors ($FCR_P^* = r_f + \delta * (g_u - g_d)$) and the transaction fees TF:

$$\frac{1}{2}H_{H,0}\underbrace{(1 + FCR_P^* + R_{P,H} + TF)}_{MR_{P,H}^*} + \frac{1}{2}[\underbrace{RR * H_{H,0}(1 + \Delta e + \Delta g + \Delta p)}_{\text{Recovery Value}} + \underbrace{I_{H,0}(1 + g_d + \Delta g - c)}_{\text{Disposable Income}}] = H_{H,0}(1 + FCR_P^* + TF) \quad (17)$$

The first term on the left-hand-side of the inequality is the product of probability in up-state ($\frac{1}{2}$) and the payment of the mortgage full amount $H_{H,0}(1 + FCR_P^* + R_{P,H} + TF)$. The second term is the product of probability in down-state ($\frac{1}{2}$) and the sum of the following two terms: (1) the recovery value of the mortgage from foreclosure sale $RR * H_{H,0}(1 + \Delta e + \Delta g + \Delta p)$ and (2) the disposable income $I_{H,0}(1 + g_d + \Delta g - c)$. The additional house price growth happens due to the additional positive growth in employment, household income, and population caused by higher net export growth. In the empirical literature, there is evidence that employment growth, household income growth, and population growth (including migration) can push up housing demand and then housing price, especially in the long run (Olsen, 1987). A perfectly competitive market means

[56] The derivation is in Appendix Section A.6.1

that the mortgage rate $MR^*_{P,H}$ is set to make the equality hold:

$$\begin{aligned} MR^*_{P,H} &:= 1 + FCR^*_P + R_{P,H} + TF \\ &= [2(1 + FCR^*_P + TF) - RR(1+\Delta e + \Delta g + \Delta p)] - (1 + g_d+\Delta g - c) * i_0 \\ &= C^*_{P,H} - K^*_{P,H} * i_0 \end{aligned} \tag{18}$$

The above mortgage rate has very clear economic interpretations. The mortgage rate in PLM in HNEG area (in a non-default state) is positively related to funding cost ($FCR^*_P = r_f + \delta * (g_u - g_d)$) and the transaction fee (TF) charged by lenders for mortgage processing and private issuers for packaging (securitization). In addition, the mortgage rate is negatively related to the recovery value of the mortgage from foreclosure sale ($RR(1+\Delta e + \Delta g + \Delta p)$) and the disposable income in the default case ($(1 + g_d+\Delta g - c) * i_0$), since these proceeds are used to cover the loss given default.

6.5 PLM in LNEG area

Similar to PLM in HNEG-are, by observing household income $I_{L,0}$ at time 0, lenders consider the household disposable income in time 1 would be $I_{L,0}(1 + g_u - \Delta g - c)$ in up-state with probability $\frac{1}{2}$ and $I_{L,0}(1 + g_d - \Delta g - c)$ in down-state with probability $\frac{1}{2}$. The full mortgage amount is $H_{L,0}(1 + FCR^*_P + R_{P,L} + TF)$. $FCR^*_P = r_f + \delta * (g_u - g_d)$ is the funding cost required by high-risk investors in equilibrium because of default risk due to a lack of insurance. $R_{P,L}$ is the premium that goes to high-risk investors in up-state to compensate for the potential loss in down-state. TF is the transaction fee rate charged by lenders for mortgage processing and by private issuers for packaging (securitization).

Assuming (1) all households can pay the full mortgage amount in up-state and (2) all households default in down-state, I get the following two boundary conditions for income-to-house ratios[57] :

$$\begin{aligned} \frac{\underline{I_{L,0}}}{H_{L,0}} &\geq \frac{1 + FCR^*_P + R_{P,L} + TF}{1 + g_u - \Delta g - c} := \underline{i_{P,L}} \\ \frac{\overline{I_{L,0}}}{H_{L,0}} &< \frac{1 + FCR^*_P + R_{P,L} + TF}{1 + g_d - \Delta g - c} := \overline{i_{P,L}} \end{aligned}$$

Therefore, I only assume that household have income-to-house ratio in time 0: $i_0 \in [\underline{i_{P,L}}, \overline{i_{P,L}})$. Again, I purposefully ignore these two extreme groups of households to make my model simple and focus my attention on the key economic mechanism of interest.

Now let us focus on the key break-even condition for the PLM market in LNEG area: the expected payoff from households shall be enough to compensate the required return of ultimate high-risk investors ($FCR^*_P = r_f + \delta * (g_u - g_d)$) and the transaction fees TF:

$$\tfrac{1}{2}H_{L,0}\underbrace{(1 + FCR^*_P + R_{P,L} + TF)}_{MR^*_{P,L}} + \tfrac{1}{2}[\underbrace{RR * H_{L,0}(1-\Delta e - \Delta g - \Delta p)}_{\text{Recovery Value}} + \underbrace{I_{L,0}(1 + g_d-\Delta g - c)}_{\text{Disposable Income}}] = H_{L,0}(1 + FCR^*_P + TF) \tag{19}$$

[57] The derivation is in Appendix Section A.6.1

The first term on the left-hand-side of the inequality is the product of probability in up-state ($\frac{1}{2}$) and the payment of the mortgage full amount $H_{L,0}(1+FCR^*_P+R_{P,L}+TF)$. The second term is the product of probability in down-state ($\frac{1}{2}$) and the sum of the following two terms: (1) the recovery value of collateral from foreclosure sale $RR*H_{L,0}(1-\Delta e-\Delta g-\Delta p)$ and (2) the disposable income $I_{L,0}(1+g_d-\Delta g-c)$. The house price growth happens due to the additional negative growth in employment, household income, and population caused by lower net export growth (essentially higher net import growth). A perfectly competitive market means that mortgage rate $MR^*_{P,L}$ is set to make the equality hold:

$$\begin{aligned} MR^*_{P,L} &\coloneqq 1+FCR^*_P+R_{P,L}+TF \\ &= [2(1+FCR^*_P+TF)-RR(1-\Delta e-\Delta g-\Delta p)]-(1+g_d-\Delta g-c)*i_0 \\ &= C^*_{P,L}-K^*_{P,L}*i_0 \end{aligned} \tag{20}$$

The above mortgage rate has very clear economic interpretations. The mortgage rate in PLM in LNEG area (in a non-default state) is positively related to funding cost ($FCR^*_P=r_f+\delta*(g_u-g_d)$) required by high-risk investors and the transaction fees (TF) charged by lenders for mortgage processing and by private issuers for packaging (securitization). In addition, the mortgage rate is negatively related to the recovery value of the mortgage from foreclosure sale ($RR(1-\Delta e-\Delta g-\Delta p)$) and the disposable income ($(1+g_d-\Delta g-c)*i_0$), since these proceeds are used to cover loss given default.

6.6 The Entire Mortgage Market: Combining GSEM and PLM

I now combine the above three cases to get a full picture of the entire mortgage market. First, I define the domain of income-to-house ratio $i_0 \in [\underline{i_0}, \overline{i_0})$ where $\underline{i_0} \coloneqq max\{\underline{i_G}, \underline{i_{P,H}}, \underline{i_{P,L}}\}$, and $\overline{i_0} \coloneqq min\{\overline{i_C}, \overline{i_{P,H}}, \overline{i_{P,L}}\}$. For the reasons stated above, my focus of $i_0 \in [\underline{i_0}, \overline{i_0})$ can help us illustrate the key economic trade-offs in a simple model.

Second, let us add lower and higher bounds of income-to-house ratios required by government-sponsored enterprise mortgages. The US government essentially subsidizes homeownership of residences by providing insurance via GSEs at a very low cost. However, such a subsidy has a lower bound criteria for household income, credit scores, income-to-mortgage ratios, and other credit factors. To incorporate these criteria into my model, I set up a corresponding lower bound of income-to-house ratio i^*_G (in both the HNEG and LNEG areas). In addition, the Federal Housing Finance Agency (FHFA) sets conforming loan limits (CLLs) for GSEM for each county each year. Mortgages above the limits are called "jumbo" mortgages and cannot be purchased by GSEs. The setup of CLLs essentially prevents the government from providing subsidies to high-income residents buying luxury houses. To incorporate this feature into my model, I set up a conforming income-to-house ratio $\overline{i_C}(<\overline{i_0})$ above which GSEs cannot purchase mortgages.

Third, let us add one ceiling for the total mortgage rates $\overline{MR_P}$ for the PLM market. The upper bound $\overline{MR_P}$ for PLM can be motivated by usury law, which is used to protect consumers from predatory lending and protect the less sophisticated consumers from irrational behavior due to poor financial knowledge.[58] Alternatively, the ceiling can be motivated by credit rationing. Mortgage lenders also have the incentive to limit PLM rates since higher rates could only attract risky borrowers (1) whose information is too limited to banks to evaluate credit risks (like many subprime mortgage borrowers) and (2) who have strong moral hazard motives. Effectively, $\overline{MR_P}$ and break-even PLM rate functions can give us a lower bound of income-to-house ratio.

6.6.1 Mortgage Market in HNEG area in normal time

Now I consider the mortgage market in HNEG area that contains both the GSEMs and PLMs in normal times when FCR_P^* is much higher than FCR_G. Based on the break-even mortgage rate functions in Equation (16) and (18), I know the absolute value of slope $K_{P,H}^*$ is higher than the one of slope K_G^* because the PLM market considers the differential higher household income growth rate Δg in HNEG area. Compared to GSEM, for each unit increase of income-to-house ratio i_0 in time 0, the PLM will be perceived as being compensated more by disposable income in default (in down-state), thereby resulting in a lower mortgage rate in equilibrium.

Now let us compare the constant part $C_{P,H}^*$ and C_G^*. For simplicity, I assume the transaction fees are the same in GSEM and PLM markets since these fees cover the same mortgage processing and packaging (securitization) process. I assume $2FCR_P^* - RR * (\Delta e + \Delta g + \Delta p) > 2FCR_G$ so that the constant part $C_{P,H}^*$ is higher than C_G^*. This assumption matches the reality in the sense that the funding cost of PLM (FCR_P^*) is much higher than the one of GSEM (FCR_G) due to a lack of insurance and potential default risk. In essence, mortgage insurance provided by GSEs is a government subsidy for home ownership so that it is charged at a very low rate GIR.[59] Therefore, the product of recovery rate and small growth rate differences in employment, household income, and population between the two areas are smaller than the twice difference between two funding costs: $2FCR_P^* - RR * (\Delta e + \Delta g + \Delta p) > 2FCR_G$. Conversely, if I assume $2FCR_P^* - RR * (\Delta e + \Delta g + \Delta p) < 2FCR_G$, together with the fact that $|K_{P,H}^* > K_G^*|$, I could get the conclusion that mortgage rate in PLM for conforming loans is even lower than mortgage rate in GSEM. This conclusion apparently contradicts reality.

The above comparison of two lines of break-even mortgage rates can be shown in Figure (4). After imposing the ceiling mortgage rate $\overline{MR_G}$ for the GSEM and $\overline{MR_P}$ for the PLM, I can solve

[58] In some states, the usury law limits are not binding due to (1) exemptions by other state laws, (2) override by contract laws, and (3) opt-out authority from federal usury laws. In these cases, I argue that the mortgage rate limits shall be interpreted broadly, including credit rationing motive.

[59] This insurance rate is lower than the economic break-even insurance rate because, in the Great Recession, the US government had to pay for defaults of a large amount of agency-backed MBS since GSEs failed to do so.

for the corresponding two lower bounds for income-to-house ratio:

$$i_G^* = \frac{2(1 + FCR_G + TF) - RR - \overline{MR_G}}{1 + g_d - c} \tag{21}$$

$$i_{P,H}^* = \frac{2(1 + FCR_P^* + TF) - RR(1+\Delta e + \Delta g + \Delta p) - \overline{MR_P}}{1 + g_d + \Delta g - c} \tag{22}$$

For easy demonstration, I further assume that $i_{P,H}^* > i_G^*$.[60] Therefore, in the mortgage lending market in HNEG area, lenders with income-to-house ratio $i_0 \in [\overline{i_C}, \overline{i_0})$ can only get mortgages from the PLM market since the government does not sponsor home ownership for the very rich group of households buying luxury houses. Lenders with income-to-house ratio $i_0 \in [i_G^*, \overline{i_C})$ would choose the GSEM market because (1) the lower rate in the GSEM market than the one in PLM market when $i_0 \in [i_{P,H}^*, \overline{i_C})$ and (2) GSEM is the only one available when $i_0 \in [i_G^*, i_{P,H}^*)$. Lastly, lenders with income-to-house ratio $i_0 \in [\underline{i_0}, i_G^*)$ cannot get mortgages because of the legal ceiling limit of PLM rates.

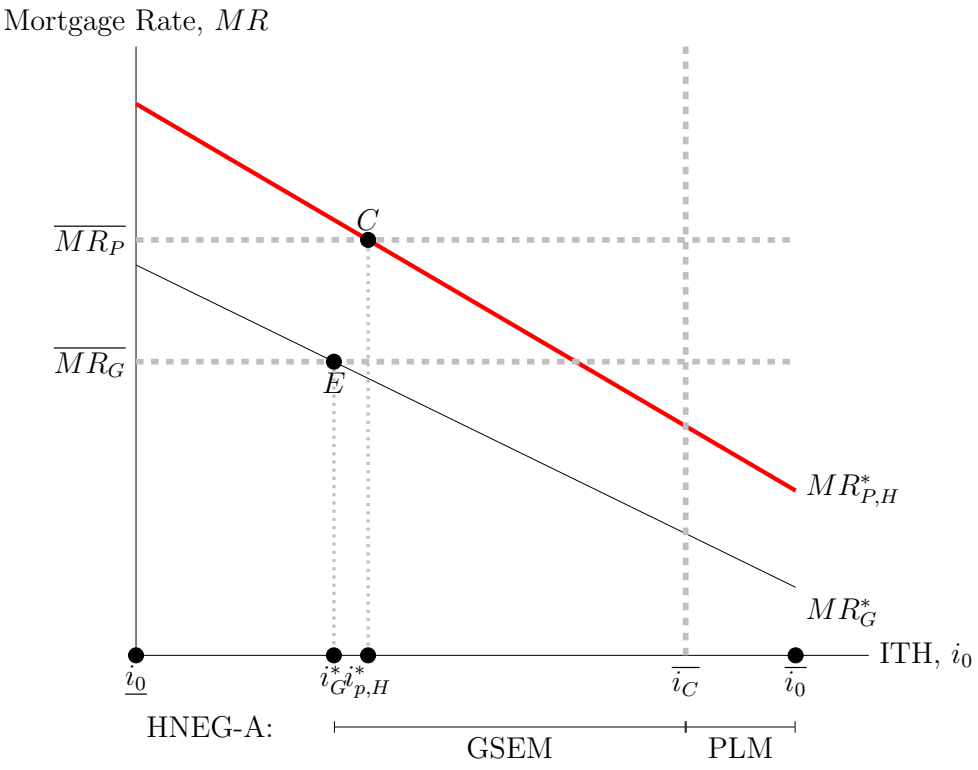


Figure 4: Mortgage Lending in High-NEG Areas

[60] This additional assumption is made to illustrate the key economic insight: the differential higher relative increase in PLM number in HNEG area than LEG-are after the decline of funding cost of PLM. By assuming $i_{P,H}^* > i_G^*$, I effectively let PLM share only cover the "jumbo" mortgages where $i_0 >= \overline{i_0}$. Of course, this assumption may not be realistic, so empirical tests of key predictions by the model are necessary.

6.6.2 Mortgage Market in LNEG area in normal time

Now I consider the mortgage market in LNEG area that contains both the GSEMs and PLMs in normal time, when FCR_P^* is much higher than FCR_G. Based on the break-even mortgage rate functions in Equation (16) and (20), I know the absolute value of slope $K_{P,L}^*$ is lower than the one of slope K_G^* because the PLM market considers the differential lower household income growth rate Δg in LNEG area. For each unit increase of income-to-house ratio i_0 in time 0, the PLM will be perceived as being compensated less by disposable income in default (in down-state), thereby charging a higher mortgage rate when there is no default (in up-state). I can also know that the constant part $C_{P,L}^*$ is higher than C_G^* because $2FCR_P^* + RR * (\Delta e + \Delta g + \Delta p) > 2FCR_G$ given that the funding cost in PLM is higher than the one in GSEM ($FCR_P^* > FCR_G$).

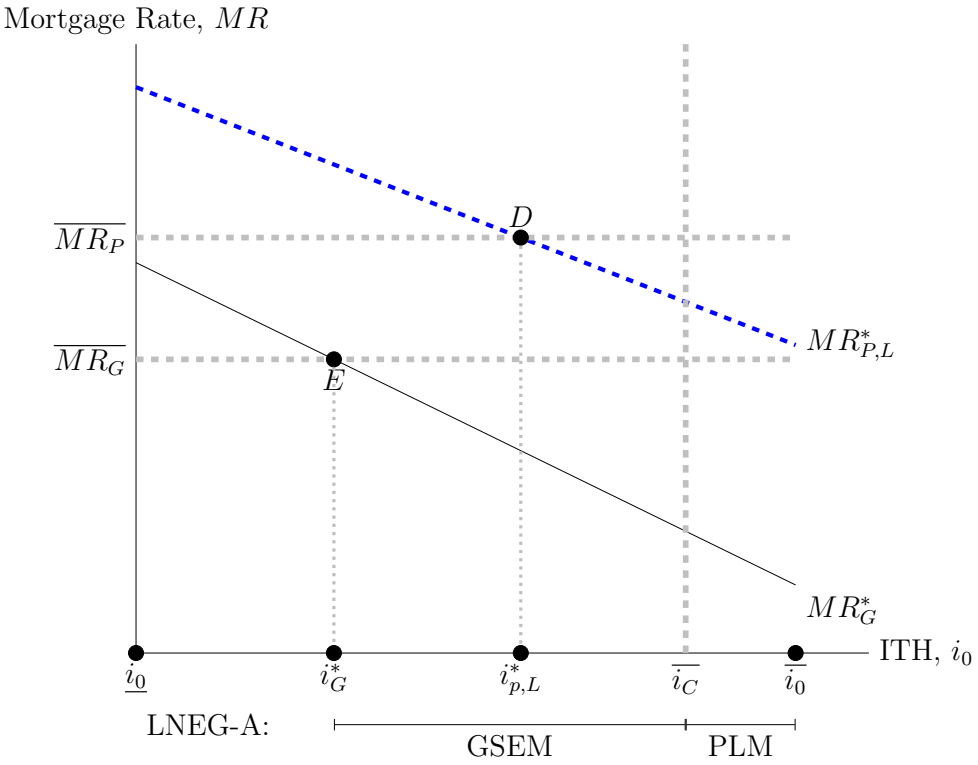


Figure 5: Mortgage Lending in Low-NEG Areas in Normal Time

The above comparison of two lines of break-even mortgage rates can be shown in Figure (5). After imposing the ceiling mortgage rate $\overline{MR_G}$ for the GSEM and $\overline{MR_P}$ for the PLM, I can additionally solve for the corresponding lower bounds for income-to-house ratio in the PLM market in LNEG area:

$$i_{P,L}^* = \frac{2(1 + FCR_P^* + TF) - RR(1 - \Delta e - \Delta g - \Delta p) - \overline{MR_P}}{1 + g_d - \Delta g - c} \tag{23}$$

By comparing numerator and denominator, I can conclude $i^*_{P,L} > i^*_{P,H}$. From previous assumption that $i^*_{P,H} > i^*_G$, I can easily know $i^*_{P,L} > i^*_G$. Therefore, in the mortgage lending market in LNEG area, lenders with income-to-house ratio $i_0 \in [\overline{i_C}, \overline{i_0})$ can only get mortgages from the PLM market since the government does not sponsor home ownership for the very rich group of households buying luxury houses. Lenders with income-to-house ratio $i_0 \in [i^*_G, \overline{i_C})$ would choose the GSEM market because (1) the lower rate in the GSEM market than the one in PLM market when $i_0 \in [i^*_{P,L}, \overline{i_C})$ and (2) GSEM is the only one available when $i_0 \in [i^*_G, i^*_{P,L})$. Lastly, lenders with income-to-house ratio $i_0 \in [\underline{i_0}, i^*_G)$ cannot get mortgages because of the legal ceiling limit of PLM rates.

6.6.3 Compare Markets in HNEG area and LNEG area in normal time

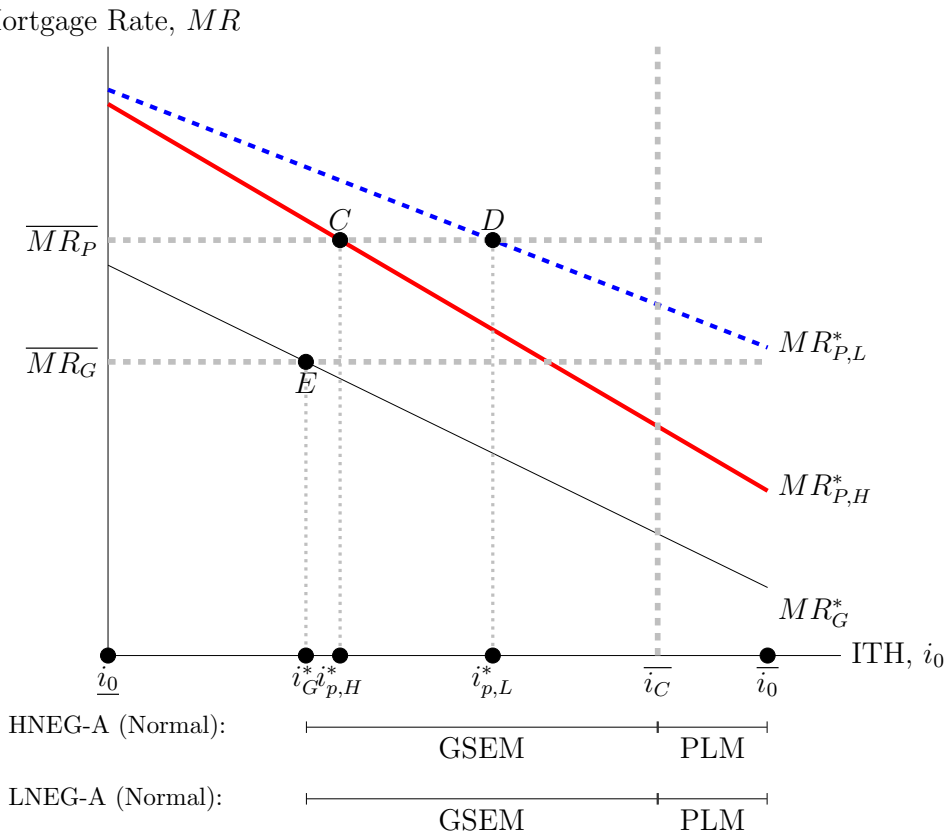


Figure 6: Mortgage Markets Before the Decline of Funding Cost in PLM Market

Now let us compare mortgage markets in HNEG area and LEG-are in normal times. Here I can compare two markets in one figure since I have assumed that income-to-house ratio i_0 follows the same distribution in both areas. First, since the GSEM market does not consider the differences in regional economic conditions, the GSEM rate is the same in both HNEG area and LNEG area.

Of course, MR^*_G is the lowest for any i_0 since GSEs provide mortgage insurance at a rate lower than the economic break-even cost, as a subsidy to home ownership by the government. Second, when $i_0 = \underline{i_0}$, I have $MR^*_{P,L} > MR^*_{P,H}$ since the PLM ultimate investors in LNEG area get lower recovery value and lower disposable income in case of default. To achieve the break-even condition, the mortgage rate must be higher for PLM in LNEG area than the one in HNEG area for the same income-to-house ratio. Third, the absolute value of the slope $K^*_{P,L} < K^*_{P,H}$, because, for each unit increase of income-to-house ratio i_0, the increase in disposable income in case of default in LNEG area is lower than the one in HNEG area. Thus, PLM rate declines with i_0 less steeply in LNEG area than in HNEG area. These patterns can be seen in Figure (6).

6.7 When Funding Cost of PLM Declines

Now I study the changes when the funding cost of PLM declines relative to the one of GSEM, the case beginning in the summer of 2003 (Justiniano, Primiceri, and Tambalotti, 2022). I can model this situation simply by letting FCR^*_P decline to FCR^{**}_P while holding FCR_G unchanged.

The above decline will result in a downward shift of PLM rates in both areas. Such shift can be seen in Figure (7). The after-shift lines are plotted in bold lines. Mortgage rate $MR^{**}_{P,H}$ and $MR^{**}_{P,L}$ interacts with legal upper limit $\overline{MR_P}$ at point A with $i_0 = i^{**}_{P,H}$ and point B with $i_0 = i^{**}_{P,L}$, respectively.

By replacing FCR^*_P with FCR^{**}_P, I can get:

$$i^{**}_{P,H} = \frac{2(1 + FCR^{**}_P + TF) - RR(1+\Delta e + \Delta g + \Delta p) - \overline{MR_P}}{1 + g_d + \Delta g - c} \tag{24}$$

$$i^{**}_{P,L} = \frac{2(1 + FCR^{**}_P + TF) - RR(1-\Delta e - \Delta g - \Delta p) - \overline{MR_P}}{1 + g_d - \Delta g - c} \tag{25}$$

By comparing both numerators and denominators, I can easily know $i^{**}_{P,H} < i^{**}_{P,L}$ so point A lies to the left of point B. I further assume $i^{**}_{P,L} < i^*_G$, which matches the reality that PLM does increase in LNEG area after its funding cost declines.[61] I also assume that $i^{**}_{P,H} > \underline{i_0}$ since, in practice, not every household in HNEG area can be approved for PLM.

[61] Please see the empirical results in this paper.

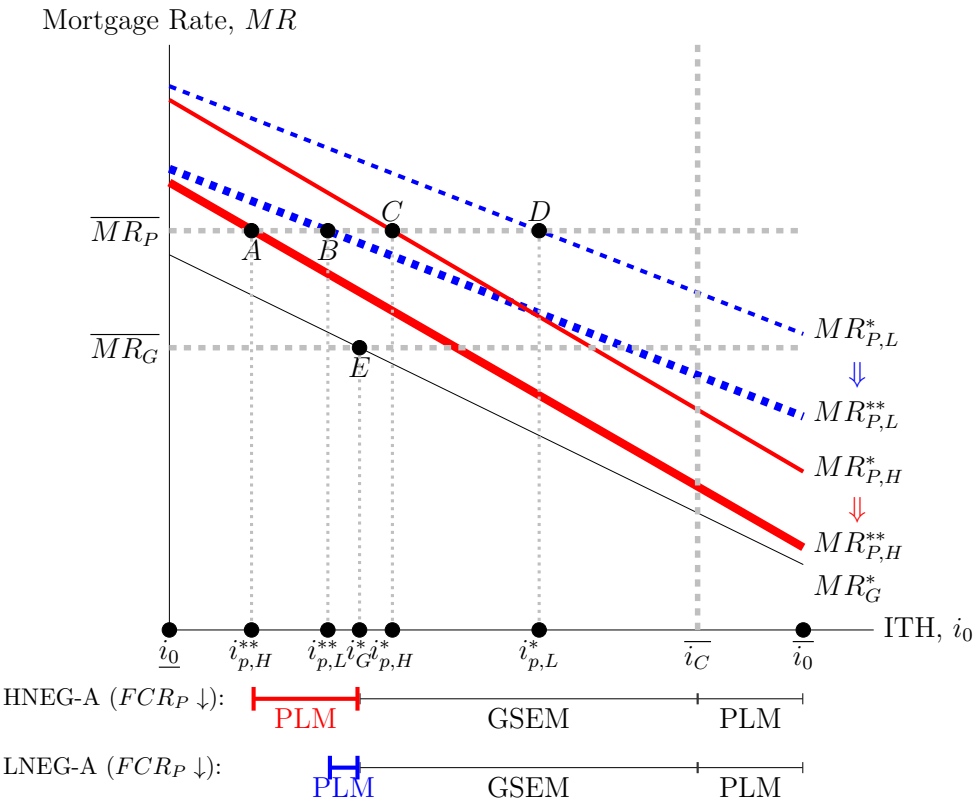


Figure 7: Mortgage Market Before and After A Decline of Funding Cost in PLM

Now, in the mortgage market in HNEG area, households with income-to-house ratio $i_0 \in [\overline{i_C}, \overline{i_0}) \cup [i^{**}_{P,H}, i^*_G)$ can only get mortgages from the PLM market. Households with income-to-house ratio $i_0 \in [i^*_G, \overline{i_C})$ would choose the GSEM market due to the lower rate. Lastly, households with income-to-house ratio $i_0 \in [\underline{i_0}, i^{**}_{P,H})$ cannot get mortgages because of the legal ceiling limit of PLM rates.

Similarly, now, in the mortgage lending market in LNEG area, households with income-to-house ratio $i_0 \in [\overline{i_C}, \overline{i_0}) \cup [i^{**}_{P,L}, i^*_G)$ can get mortgages from the PLM market. Households with income-to-house ratio $i_0 \in [i^*_G, \overline{i_C})$ would choose the GSEM market due to the lower rate. Lastly, households with income-to-house ratio $i_0 \in [\underline{i_0}, i^{**}_{P,L})$ cannot get mortgages because of the legal ceiling limit of PLM rates.

6.8 Summary: How the Model Reconciles Two Empirical Facts

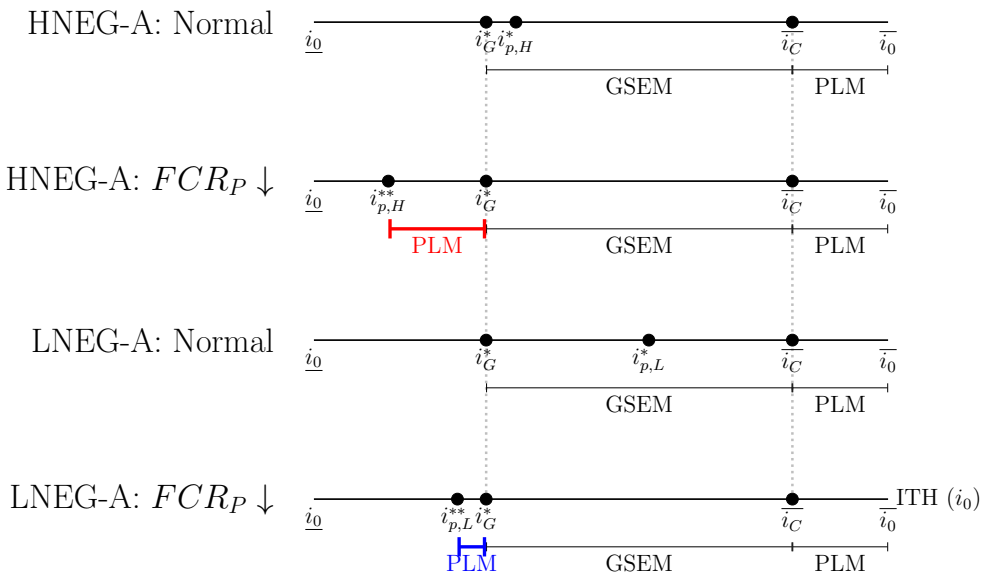


Figure 8: Compare: Normal Time vs FCR_P^* Declines

Figure (8) summarizes the mortgage market shares by GSEM and PLM both in normal times (when the funding cost of PLM (FCR_P^*) is much higher than the one of GSEM (FCR_G)) and when FCR_P^* declines. In the latter case, my model predicts that credit expansion in PLM is stronger in the HNEG area than in the LNEG area. This stronger increase in PLM share in the HNEG area shows up in the relationship $i_{P,H}^{**} < i_{P,L}^{**}$. Comparing Equation (24) and (25), a sufficient condition is that $\Delta g > 0$, $\Delta e > 0$, and $\Delta p > 0$. In economic terms, this sufficient condition means there are higher positive growth rates in employment, household income, and population in the high-net-export-growth areas in the future (unrealized prediction).[62] It is through these three economic mechanisms (unrealized prediction) that net export growth drives the stronger credit supply in PLM in the HNEG area between 1999-2005. For empirical tests verifying that higher net export growth causes higher employment growth, higher household income growth, and higher population growth, see Appendix Section A.5.

The prediction that credit expansion in PLM is stronger in the HNEG area implies that the positive correlation between mortgage growth and income growth across metropolitan areas is a natural result of credit expansion in response to net export growth. This phenomenon happens because (1) within each metro, ZIP codes with low income-to-house ratio experience the majority

[62] As argued in Section 5, realized effects at the household level could be considered by government-sponsored enterprise mortgages, which seems not big enough to show up in cross-metro patterns of GSEMs. In contrast, unrealized and predicted effects at the metropolitan level can only be considered by private-label mortgages, which is large enough to show up in tests in Section 4.

of PLM growth, and (2) ZIP codes with low income-to-house ratio in the HNEG areas experience much higher growth in PLMs than the ones in the LNEG areas. This empirical fact alone is not sufficient to support the "speculation" view as Adelino, Schoar, and Severino (2016). Therefore, the empirical evidence provided in support of this model argues against the "speculation" interpretation by Adelino, Schoar, and Severino (2016).

The model also predicts that a natural result of credit expansion is that within-metropolitan areas (and within-county), credit expansion is stronger in low income-to-house-ratio areas. This prediction is consistent with the empirical evidence by Mian and Sufi (2009) that subprime zip codes experience stronger mortgage growth within counties.

Thus, the simple model reconciles two seemly opposing empirical facts: the correlation between income growth and mortgage growth is positive across metropolitan areas (Adelino, Schoar, and Severino, 2016) but negative areas across ZIP codes within metropolitan areas (Mian and Sufi, 2009). These two empirical facts can coexist within a credit expansion framework.

7 Model Predictions of "Double Differences" and Causal Evidence

Beyond reconciling two seemingly opposing empirical facts within the "credit expansion" view, this model provides new predictions of "double differences" in mortgage growth and house price growth. This section develops these new predictions and provides causal evidence and robustness checks.

7.1 Model-based New Predictions of "Double Differences"

Hypothesis 4 The 1999-2005 difference in mortgage growth between the low-income ZIP codes and the high-income ZIP codes within the same metropolitan areas is positively caused by net export growth (1999-2005) across metropolitan areas. However, the corresponding 2005-2008 difference is negatively caused by net export growth (1999-2005) across metropolitan areas.

This hypothesis starts from the basic prediction that most credit expansion in mortgage growth concentrates in the low-income zip codes. In the argument by Mian and Sufi (2009), an implicit assumption (and a fact in data) is that households choose ZIP codes within metropolitan areas primarily based on their income-to-house ratio. In my model, this assumption means households with a low income-to-house ratio primarily live in low-income ZIP codes, while households with a high income-to-house ratio live in high-income ZIP codes. I adopt this assumption and create a differential mortgage growth between low-income ZIP codes and high-income ones within metropolitan areas: "low-minus-high" differential mortgage growth. Due to credit expansion in response to net export growth, I predict that this "low-minus-high" differential mortgage growth is positively correlated with net export growth (99-05) between 1999 and 2005 but negatively between 2005 and

2008.[63] This prediction can be seen intuitively from Figure 8. Within metropolitan areas, credit expansion in private-label mortgages primarily centers on the low-income ZIP codes (left tail). Thus, the 'low-minus-high" differential mortgage growth primarily captures the mortgage growth in the low-income ZIP codes. Due to the household advantages caused by net export growth, the 'low-minus-high" differential mortgage growth is much higher in the high net-export-growth metropolitan areas than in the low net-export-growth ones. In addition, since the credit expansion is unsustainable, the subsequent reversal is negative in magnitude.

Hypothesis 5 The 2000-2006 difference in house price growth between the low-income ZIP codes and the high-income ZIP codes within the same metropolitan areas is positively caused by "low-minus-high" differential mortgage growth (1999-2005) across metropolitan areas. However, the corresponding 2007-2009 difference in house price growth is negatively caused by "low-minus-high" differential mortgage growth (1999-2005) across metropolitan areas.

The above hypothesis is derived by applying the same logic to differential house price growth. Though Mian and Sufi (2009) and Adelino, Schoar, and Severino (2016) also use counties, metropolitan areas are better units to study mortgage and housing markets across ZIP codes within an area of similar economic condition. First, counties are smaller than metropolitan areas. Some counties are in the central or richer places while others are on the outskirts or poor in metropolitan areas. Thus, the cross-ZIP code variation within metropolitan areas is larger than within counties. Second, due to the frequent commuting across counties within a metropolitan area, the net export growth measure is more appropriately aggregated at the metropolitan level than at the county level.

7.2 "Double differences" in Mortgage Growth: Causal Evidence

This subsection provides descriptive evidence in figures and causal evidence in regressions for the model-based new prediction in differential mortgage growth (hypotheses 4).

First, Figure (14) depicts the private-label mortgage (non-jumbo) growth (92-09) across ZIP code groups (high vs. low income) and across metropolitan areas (high vs. low net-export-growth). In Subfigure (a) with top and bottom income half ZIP codes, in the prior period (92-99), differential mortgage growth between high vs. low-income ZIP codes within metropolitan areas is very close to zero. In addition, differential mortgage growth between high vs. low net-export-growth metropolitan areas is also very close to zero. However, in the boom period (99-05), differential mortgage growth between high vs. low-income ZIP codes within metropolitan areas (hereafter "high-minus-low" mortgage growth) is large. In addition, this "high-minus-low" mortgage growth is larger in the high net-export-growth metropolitan areas than the one in the low net-export-growth ones.

[63] This model-based new prediction is not a natural result by combining two empirical facts by Adelino, Schoar, and Severino (2016); Mian and Sufi (2009). It could be the case that "low-minus-high" differential mortgage growth is the same across metropolitan areas.

Subfigure (b) shows similar trends with top and bottom quintile-income ZIP codes. However, since quintile groups compare two end tails, the differential "high-minus-low" mortgage growth across metropolitan areas seems even more pronounced in Subfigure (b) than in Subfigure (a).

To formally test the new prediction, we use the following regression specification:

$$\begin{aligned}\text{LMH}\triangle_{99,05}\&\triangle_{05,08} Ln(PLMNJ_m) = \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \beta_{05,08} * \triangle_{99,05}\text{NetExp}_m \times Dum_{05,08} \\ + \gamma_{99,05} * Dum_{99,05} + \gamma_{05,08} * Dum_{05,08} + \epsilon_{period,m}\end{aligned} \tag{26}$$

where $\triangle_{99,05}\text{givNetExp}_m$ is the IV for $\triangle_{99,05}\text{NetExp}_m$.

Table (19) reports the regression results for OLS, reduced-form, first-stage, and second-stage results. Four columns report the above four regressions for the sample of top and bottom of half, tertile, quartile, and quintile income groups ZIP codes, respectively. Two clear patterns show up. First, when ZIP codes are divided into finer groups based on 1998 household income level, the results of "double differences" are more striking in magnitude. The OLS, reduced-form, and 2SLS estimates in the boom period (99-05) are not significant in the samples of half and tertile income groups but statistically significant in the samples of quartile and quintile income groups. The coefficient equality tests fail to reject equality in columns (1) and (2), but reject equality in columns (3) and (4). Likewise, the adjusted R-squares in OLS and reduced-form regressions are gradually getting larger from column (1) to column (3), with column (4) being very close to column (3). Second, the results in the bust period (05-08) are more pronounced than in the boom period (99-05). The reduced-form and 2SLS estimates in the bust period (05-08) are not significant in the samples of half-income groups but significant in the samples of tertile, quartile, and quintile income groups. OLS estimates are significant in the bust period in all columns, while estimates are only significant in the boom period in columns (3) and (4). In summary, the empirical tests show that net export growth (99-05) induces a much stronger "low-minus-high" differential mortgage boom (99-05) and bust (05-08) cycle in the high net-export-growth areas.

7.3 "Double differences" in House Price Growth: Causal Evidence

Figure (20) depicts the house price growth (92-09) across ZIP code groups (high vs. low income) across metropolitan areas (high vs. low net-export-growth). In Subfigure (a) with top and bottom half-income ZIP codes, in the prior period (92-99), differential house price growth between high vs. low-income ZIP codes within metropolitan areas is very close to zero. In addition, differential house price growth between high vs. low net-export-growth metropolitan areas is also very close to zero. However, in the boom period (00-06), differential house price growth between high vs. low-income ZIP codes within metropolitan areas (hereafter "high-minus-low" house price growth) is large.[64] In addition, this "high-minus-low" house price growth is larger in the high net-export-

[64] Unlike analysis across metropolitan areas, We use 2000-2006 as the boom period for house price growth across ZIP codes within metropolitan areas because the ZIP-level double difference analysis (hypothesis 5) requires more

growth metropolitan areas thae in the low net-export-growth ones. Subfigure (b) shows similar trends with top and bottom quintile-income ZIP codes. However, since quintile groups compare two end tails, the differential "high-minus-low" house price boom and bust across metropolitan areas is even more pronounced in Subfigure (b) than in Subfigure (a).

To formally test the new prediction, we use the following regression specification:

$$\begin{aligned}\text{LMH}\triangle_{00,06}\&\text{LMH}\triangle_{07,09}Ln(HPI_m) = \beta_{00,06} * \text{LMH}\triangle_{99,05}Ln(PLMNJ_m) \times Dum_{00,06} \\ + \beta_{07,09} * \text{LMH}\triangle_{99,05}Ln(PLMNJ_m) \times Dum_{07,09} + \alpha_{00,06} + \alpha_{07,09} + \epsilon_{period,m}\end{aligned} \tag{27}$$

where $\triangle_{99,05}\text{givNetExp}_m$ is the IV for $\triangle_{99,05}\text{NetExp}_m$. Table (20) reports the regression results for OLS, reduced-form, first-stage, and second-stage results. Four columns report the above four regressions for the sample of top and bottom of half, tertile, quartile, and quintile groups ZIP codes, respectively. Two clear patterns are apparent. First, when ZIP codes are further divided into finer groups based on 1998 household income, the empirical results of "double differences" are more striking. The 2SLS estimates in the boom period (00-06) are insignificant in the samples of half and tertile income groups but significant in the samples of quartile and quintile income groups. The coefficient equality tests reject equality in columns (3) and (4). Likewise, the coefficient magnitudes in OLS and reduced-form regressions are gradually getting larger from column (1) to column (3), whereas column (4) is very close to column (3). Second, the results in the bust period (05-08) are more pronounced than in the boom period (00-06). The coefficients of OLS, reduced-form, and 2SLS estimates in the bust period (07-09) are around four to seven times the ones in the boom period (00-06). In summary, the empirical tests show that induced by net export growth (99-05), the "low-minus-high" differential mortgage growth causes a much stronger "low-minus-high" differential house price boom (00-06) and bust (07-09) cycle in the high net-export-growth areas.

7.4 "Double differences": Robustness

This subsection provides robustness tests of model-based predictions of "double differences" using the number of private-label mortgages (non-jumbo) and dollar amount of owner-occupied private-label mortgages (non-jumbo).

First, we show that the model-based prediction of "double differences" in mortgage growth applies to the number of private-label mortgages (non-jumbo). The number of loans is also important because it represents the volume (or churning) of transactions in the housing market. Figure (15) depicts the private-label mortgage (non-jumbo) number growth (92-09) across ZIP code groups (high vs. low income) and across metropolitan areas (high vs. low net-export-growth). Similar to the patterns of the dollar amount of private-label mortgages (non-jumbo), Subfigure (a) and (b)

data variation, and house price growth shows a one-year lag in the data. This one-year lag (or even two-year lag) can be seen in Figure (14) and (20). The differential "high-minus-low" mortgage growth across metropolitan areas is the largest ending in 2005. However, the differential "high-minus-low" house price boom across metropolitan areas is the largest ending in 2006 (or even 2007).

show that in the prior period (92-99) "high-minus-low" mortgage growth is very close to zero, and its difference between high vs. low net-export-growth metro areas is also very close to zero. However, in the boom period (99-05), "high-minus-low" mortgage growth is large in absolute number and is larger in the high net-export-growth areas. Again, the above "double difference" pattern seems to be stronger in the Subfigure (b) with finer-divided zip codes based on income level.

For the regression tests, we perform a similar specification in equation (26) except that the dependent variable is the "low-minus-high" differential mortgage number growth. Table (21) reports the regression results for OLS, reduced-form, first-stage, and second-stage results. As before, the same two patterns are apparent. First, when ZIP codes are divided into finer groups based on 1998 household income, the results of "double differences" are more striking. Second, the estimates in the bust period (05-08) are much stronger than the ones in the boom period (99-05). In general, the regressions show that net export growth (99-05) induces a much stronger "low-minus-high" differential mortgage boom (99-05) and bust (05-08) cycle in number of loans in the high net-export-growth areas.

Second, we show that the model-based prediction of "double differences" in differential mortgage growth applies to the dollar amount of owner-occupied private-label mortgages (non-jumbo). which is used as a measure pure credit expansion in Section 5. The goal of this robustness is to show that the "double differences" pattern is independent of potential speculation. Figure (16) depicts the owner-occupied private-label mortgage (non-jumbo) dollar growth (92-09) across ZIP code groups (high vs. low income) and across metropolitan areas (high vs. low net-export-growth). Similar to the results before, Subfigure (a) and (b) show that in the prior period (92-99) "high-minus-low" owner-occupied mortgage growth is very close to zero, and its difference between high vs. low net-export-growth metro areas is also very close to zero. However, in the boom period (99-05), "high-minus-low" owner-occupied mortgage growth is large in absolute magnitude and is larger in the high net-export-growth areas. Again, the above "double difference" pattern appears to be stronger in the Subfigure (b) with finer-divided zip codes.

For the regressions, we employ the specification in equation (26) except that the dependent variable is the "low-minus-high" owner-occupied mortgage growth. Table (22) reports the regression results for OLS, reduced-form, first-stage, and second-stage results. As before, the same two patterns are apparent. First, when ZIP codes are divided into more groups based on 1998 household income, the patterns of "double differences" are more striking. Second, the estimates in the bust period (05-08) are much larger in magnitude than the ones in the boom period (99-05). In summary, the regressions show that net export growth (99-05) induces a much stronger "low-minus-high" owner-occupied mortgage boom (99-05) and bust (05-08) cycle in the high net-export-growth areas.

Third, we provide evidence that the model-based predictions of "double differences" are not

driven by demand or speculation motive in borrowers (households). Again, we use the distinction that government-sponsored enterprise mortgages require a certain level of household credit ability. If demand or speculation were to explain the above results of "double differences", we shall expect that households with qualified credit ability would also demand or speculate on government-sponsored enterprise mortgages (GSEMs) such that "double differences" show up in GSEMs. However, empirical evidence shows that "double differences" do not show up in GSEMs, thus arguing against the demand or speculation view.

Figure (17) depicts the government-sponsored enterprise mortgage growth (92-09) across ZIP code groups (high vs. low income) and across metropolitan areas (high vs. low net-export-growth). Unlike the previous patterns of private-label mortgages (non-jumbo), Subfigure (a) and (b) show that in the prior period (92-99) and the boom period (99-05), "high-minus-low" mortgage growth is not substantially positive. Most importantly, its difference between high vs. low net-export-growth metro areas is not substantially different zero. Further, the above "double difference" pattern seems not to be stronger in the Subfigure (b) with finer-divided zip codes based on income level.

For the regression tests, we perform a similar specification in equation (26) except that the dependent variable is the "low-minus-high" differential growth rate in dollar amount (07USD) in government-sponsored enterprise mortgages. Table (23) reports the regression results for OLS, reduced-form, first-stage, and second-stage results. Unlike previous results, two previous patterns do not show up. First, when ZIP codes are divided into finer groups based on 1998 household income, the results of "double differences" are not more significant. Second, the estimates in the bust period (05-08) are not stronger than the ones in the boom period (99-05). In general, the regressions show that net export growth (99-05) does not induce a stronger "low-minus-high" differential boom (99-05) and bust (05-08) cycle in government-sponsored enterprise mortgages in the high net-export-growth areas.

8 Conclusion

The U.S. housing boom and bust in the 2000s were unprecedented, resulting in the deepest recession since the Great Depression. There is continuing debate regarding the cause between the "supply-side credit expansion" view vs. the "demand-side speculation" view. These two views are based on two seemingly opposing empirical facts: the correlation between income growth and mortgage growth is negative within metropolitan areas (which supports the "credit expansion" view) but positive across metropolitan areas (which supports the "speculation" view).

This paper reconciles the above two empirical facts within the "credit expansion" view with empirical causal evidence and with a theoretical model. My core finding is that credit expansion, rather than speculation, is dominant in causing the 1999-2009 U.S. housing cycle. Therefore, the

policy implies that the new regulatory design to prevent the same housing cycle shall focus more on the lenders.

In the first part of my paper, I show that the "speculation" interpretation regarding the positive correlation between income growth and mortgage growth across metropolitan areas is potentially misguided due to an omitted variable problem. The "speculation" view omits regional economic growth, for which I proxy with metropolitan exposure to net export growth of manufacturing industries (hereafter "net export growth"). Using an instrumental variable approach from international economics, I show that net export growth across metropolitan areas causes income growth and credit expansion in mortgage growth, which eventually leads to the housing cycle. I also design five tests to show that credit expansion rather than speculation dominates this housing cycle. In the second part, I use a model to show that credit expansion can reconcile the above two empirical facts, thus addressing the potential concern regarding the internal conflicts between the two seemingly opposing empirical facts. This model generates new predictions of "double differences": the differential stronger boom and bust cycle in mortgages (and house prices) in the low-income ZIP codes than the high-income ZIP codes within the same metropolitan areas positively correlated with net export growth across metropolitan areas. I find empirical causal evidence for these model-based new predictions of "double differences", further supporting the "credit expansion" view.

9 References

Acemoglu, Daron, David Autor, David Dorn, Gordon H Hanson, and Brendan Price, 2016, Import competition and the great us employment sag of the 2000s, *Journal of Labor Economics* 34, S141–S198.

Adelino, Manuel, Antoinette Schoar, and Felipe Severino, 2012, Credit supply and house prices: evidence from mortgage market segmentation, Discussion paper, National Bureau of Economic Research.

——— , 2016, Loan originations and defaults in the mortgage crisis: The role of the middle class, *The Review of Financial Studies* 29, 1635–1670.

Amiti, Mary, and Caroline Freund, 2010, The anatomy of china's export growth, in *China's growing role in world trade* . pp. 35–56 (University of Chicago Press).

Anenberg, Elliot, Aurel Hizmo, Edward Kung, and Raven Molloy, 2019, Measuring mortgage credit availability: A frontier estimation approach, *Journal of Applied Econometrics* 34, 865–882.

Autor, David, David Dorn, and Gordon H Hanson, 2013a, The china syndrome: Local labor market effects of import competition in the united states, *American Economic Review* 103, 2121–68.

Autor, David H, David Dorn, and Gordon H Hanson, 2013b, The china syndrome: Local labor market effects of import competition in the united states, *American economic review* 103, 2121–2168.

Avery, Robert, Kenneth Brevoort, and Glenn Canner, 2007, Opportunities and issues in using hmda data, *Journal of Real Estate Research* 29, 351–380.

Avery, Robert B, Neil Bhutta, Kenneth P Brevoort, Glenn B Canner, Christa N Gibbs, et al., 2010, The 2008 hmda data: the mortgage market during a turbulent year, *Federal Reserve Bulletin* 96, 169–211.

Bailey, Amy Kate, and Allison Suppan Helmuth, 2023, National zip code crosswalk (1990-2020), Discussion paper, Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor].

Barrot, Jean-Noël, Erik Loualiche, Matthew Plosser, and Julien Sauvagnat, 2022, Import competition and household debt, *The Journal of Finance* 77, 3037–3091.

Bernanke, Ben, 2005, The global saving glut and the us current account deficit, Discussion paper, Board of Governors of the Federal Reserve System (US).

——— , et al., 2007, Global imbalances: recent developments and prospects, *Bundesbank Lecture speech, September* 4, 18.

Bogin, Alexander N, William M Doerner, and William D Larson, 2019, Missing the mark: Mortgage valuation accuracy and credit modeling, *Financial Analysts Journal* 75, 32–47.

Broadman, Harry G, 2006, *From disintegration to reintegration: Eastern Europe and the former Soviet Union in international trade* (World Bank Publications).

Choi, Hyun-Soo, Harrison G Hong, Jeffrey D Kubik, and Jeffrey P Thompson, 2016, Sand states and the us housing crisis, *Available at SSRN 2373179*.

Dell'Ariccia, Giovanni, Deniz Igan, and Luc UC Laeven, 2012, Credit booms and lending standards: Evidence from the subprime mortgage market, *Journal of Money, Credit and Banking* 44, 367–384.

Deryugina, Tatyana, Laura Kawano, and Steven Levitt, 2018, The economic impact of hurricane katrina on its victims: Evidence from individual tax returns, *American Economic Journal: Applied Economics* 10, 202–233.

Di Maggio, Marco, and Amir Kermani, 2017, Credit-induced boom and bust, *The Review of Financial Studies* 30, 3711–3758.

Favara, Giovanni, and Jean Imbs, 2015, Credit supply and the price of housing, *American Economic Review* 105, 958–992.

Feenstra, Robert C, Robert E Lipsey, Haiyan Deng, Alyson Ma, and Hengyong Mo, 2005, World trade flows: 1962-2000, .

Feenstra, Robert C, Hong Ma, and Yuan Xu, 2019, Us exports and employment, *Journal of International Economics* 120, 46–58.

Feenstra, Robert C, and John Romalis, 2014, International prices and endogenous quality, *The Quarterly Journal of Economics* 129, 477–527.

Ganduri, Rohan, 2023, What drives screening incentives in nonbank mortgage originators?, *Real Estate Economics* 51, 1321–1355.

Gao, Zhenyu, Michael Sockin, and Wei Xiong, 2020, Economic consequences of housing speculation, *The Review of Financial Studies* 33, 5248–5287.

Ghent, Andra C, and Marianna Kudlyak, 2011, Recourse and residential mortgage default: evidence from us states, *The Review of Financial Studies* 24, 3139–3186.

Greenland, Andrew, John Lopresti, and Peter McHenry, 2019, Import competition and internal migration, *Review of Economics and Statistics* 101, 44–59.

Greenwood, Michael J, 1997, Internal migration in developed countries, in Mark R. Rosenzweig, and Oded Stark, ed.: *Handbook of population and family economics*vol. 1 . chap. 12, pp. 647–720 (Elsevier).

Griffin, John M, Samuel Kruger, and Gonzalo Maturana, 2021, What drove the 2003–2006 house price boom and subsequent collapse? disentangling competing explanations, *Journal of Financial Economics* 141, 1007–1035.

Griffin, John M, and Gonzalo Maturana, 2016a, Did dubious mortgage origination practices distort house prices?, *The Review of Financial Studies* 29, 1671–1708.

———, 2016b, Who facilitated misreporting in securitized loans?, *The Review of Financial Studies* 29, 384–419.

Holmes, Thomas J, and John J Stevens, 2004, Spatial distribution of economic activities in north america, in Vern Henderson, and Jacques-François Thisse, ed.: *Handbook of regional and urban economics*vol. 4 . pp. 2797–2843 (Elsevier).

Hoynes, Hilary, Douglas L Miller, and Jessamyn Schaller, 2012, Who suffers during recessions?, *Journal of Economic perspectives* 26, 27–48.

Hurst, Erik, Benjamin J Keys, Amit Seru, and Joseph Vavra, 2016, Regional redistribution through the us mortgage market, *American Economic Review* 106, 2982–3028.

Imbens, Guido W, and Joshua D Angrist, 1994, Identification and estimation of local average treatment effects, *Econometrica: journal of the Econometric Society* pp. 467–475.

Jiang, Wei, 2017, Have instrumental variables brought us closer to the truth, *The Review of Corporate Finance Studies* 6, 127–140.

Justiniano, Alejandro, Giorgio E Primiceri, and Andrea Tambalotti, 2022, The mortgage rate conundrum, *Journal of Political Economy* 130, 121–156.

Kaplan, Greg, Kurt Mitman, and Giovanni L Violante, 2020, Non-durable consumption and housing net worth in the great recession: Evidence from easily accessible data, *Journal of Public Economics* 189, 104176.

Keys, Benjamin J, Tanmoy Mukherjee, Amit Seru, and Vikrant Vig, 2010, Did securitization lead to lax screening? evidence from subprime loans, *The Quarterly journal of economics* 125, 307–362.

Keys, Benjamin J, Tomasz Piskorski, Amit Seru, and Vikrant Vig, 2013, Mortgage financing in the housing boom and bust, in L. Glaeser Edward, and Sinai Todd, ed.: *Housing and the Financial Crisis* . chap. 4, pp. 143–204 (University of Chicago Press).

Kindleberger, Charles P, 1978, Manias, panics, and rationality, *Eastern Economic Journal* 4, 103–112.

Kleibergen, Frank, and Richard Paap, 2006, Generalized reduced rank tests using the singular value decomposition, *Journal of econometrics* 133, 97–126.

Krugman, Paul, 1991, Increasing returns and economic geography, *Journal of political economy* 99, 483–499.

——— , Maurice Obstfeld, and Marc Melitz, 2018, *Chapter 7. External Economies of Scale And the International Location of Production* (Pearson).

Krugman, Paul R, 2008, Trade and wages, reconsidered, *Brookings papers on economic activity* 2008, 103–154.

Kwon, Spencer Yongwook, Yueran Ma, and Kaspar Zimmermann, 2022, 100 years of rising corporate concentration, .

Lewis, Brittany Almquist, 2023, Creditor rights, collateral reuse, and credit supply, *Journal of Financial Economics* 149, 451–472.

Li, David X, 2000, A copula function approach, *Journal of Fixed Income* 9, 43–54.

Li, Xue, and Zehan Pan, 2023, Globalization-driven internal migration in china: The impact of foreign direct investment and exports since 2000, *Population, Space and Place* 29, e2612.

Ling, David C, and Wayne R Archer, 2017, Market determinants of value, in *Real estate principles: A value approach* . chap. 5, pp. 99–128 (McGraw-Hill Education).

Mayer, Christopher, Karen Pence, and Shane M Sherlund, 2009, The rise in mortgage defaults, *Journal of Economic perspectives* 23, 27–50.

Mian, Atif, Kamalesh Rao, and Amir Sufi, 2013, Household balance sheets, consumption, and the economic slump, *The Quarterly Journal of Economics* 128, 1687–1726.

Mian, Atif, and Amir Sufi, 2009, The consequences of mortgage credit expansion: Evidence from the us mortgage default crisis, *The Quarterly journal of economics* 124, 1449–1496.

——— , 2011, House prices, home equity–based borrowing, and the us household leverage crisis, *American Economic Review* 101, 2132–2156.

——— , 2014, What explains the 2007–2009 drop in employment?, *Econometrica* 82, 2197–2223.

——— , 2017, Fraudulent income overstatement on mortgage applications during the credit expansion of 2002 to 2005, *The Review of Financial Studies* 30, 1832–1864.

——— , 2022, Credit supply and housing speculation, *The Review of Financial Studies* 35, 680–719.

——— , and Francesco Trebbi, 2013, The political economy of the subprime mortgage credit expansion, *Quarterly Journal of Political Science* 8, 373–408.

——— , 2015, Foreclosures, house prices, and the real economy, *The Journal of Finance* 70, 2587–2634.

Minot, Nicholas, and Francesco Goletti, 1998, Export liberalization and household welfare: the case of rice in vietnam, *American Journal of Agricultural Economics* 80, 738–749.

Minsky, H, 1986, Stabilizingan unstable economy yale university press, *New Haven, Connecticut*.

Molloy, Raven, Christopher L Smith, and Abigail Wozniak, 2011, Internal migration in the united states, *Journal of Economic perspectives* 25, 173–96.

Nijkamp, Peter, Piet Rietveld, and Folke Snickars, 1987, Regional and multiregional economic models: a survey, vol. 1 . chap. 7, pp. 257–294 (Elsevier).

Olea, José Luis Montiel, and Carolin Pflueger, 2013, A robust test for weak instruments, *Journal of Business & Economic Statistics* 31, 358–369.

Olsen, Edgar O, 1987, The demand and supply of housing service: a critical survey of the empirical literature, in *Handbook of regional and urban economics*vol. 2 . pp. 989–1022 (Elsevier).

Pierce, Justin R, and Peter K Schott, 2016, The surprisingly swift decline of us manufacturing employment, *American Economic Review* 106, 1632–1662.

Piskorski, Tomasz, Amit Seru, and James Witkin, 2015, Asset quality misrepresentation by financial intermediaries: Evidence from the rmbs market, *The Journal of Finance* 70, 2635–2678.

Purnanandam, Amiyatosh, 2011, Originate-to-distribute model and the subprime mortgage crisis, *The review of financial studies* 24, 1881–1915.

Redding, Stephen J, and Matthew A Turner, 2015, Transportation costs and the spatial organization of economic activity, in Gilles Duranton, Vernon Henderson, and William Strange, ed.: *Handbook of regional and urban economics*vol. 5 . chap. 20, pp. 1339–1398 (Elsevier).

Rodrik, Dani, and Arvind Subramanian, 2004, From" hindu growth" to productivity surge: the mystery of the indian growth transition, .

Romalis, John, 2007, Nafta's and cusfta's impact on international trade, *The Review of Economics and Statistics* 89, 416–435.

Saiz, Albert, 2010, The geographic determinants of housing supply, *The Quarterly Journal of Economics* 125, 1253–1296.

Salmon, Felix, 2012, The formula that killed wall street, *Significance* 9, 16–20.

Sanderson, Eleanor, and Frank Windmeijer, 2016, A weak instrument f-test in linear iv models with multiple endogenous variables, *Journal of econometrics* 190, 212–221.

Sherlund, Shane M, 2008, The jumbo-conforming spread: A semiparametric approach, .

Sykes, Jay B, 2019, Federal preemption in the dual banking system: An overview and issues for the 116th congress, *Congressional Research Service Report*.

The World Bank, 2009, *Chapter 4. Scale Economies and Agglomeration* (The World Bank).

Thrall, Grant Ian, 2002, Understanding real estate markets and submarkets, in *Business geography and new real estate market analysis* . chap. 2, pp. 19–59 (Oxford University Press, USA).

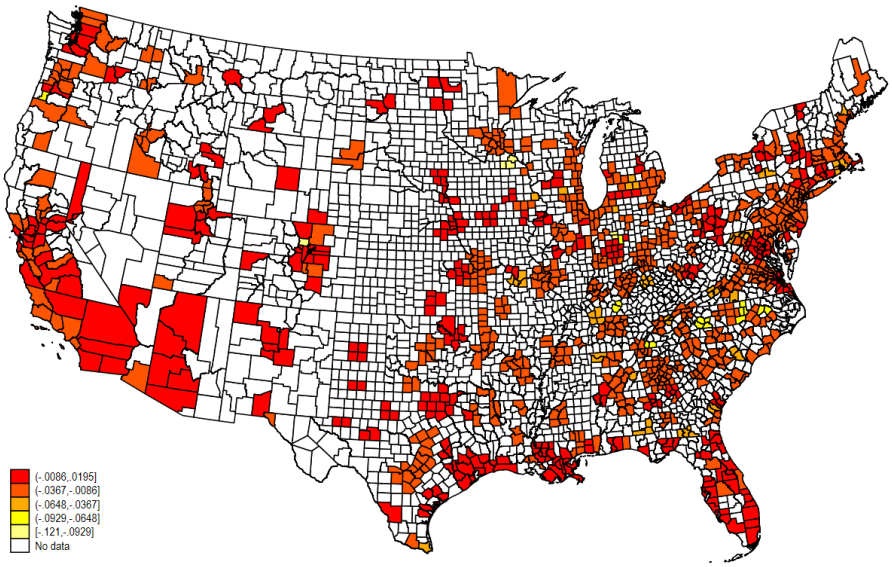
Tiebout, Charles M, 1956, A pure theory of local expenditures, *Journal of political economy* 64, 416–424.

——— , 1962, The community economic base study, Discussion paper, .

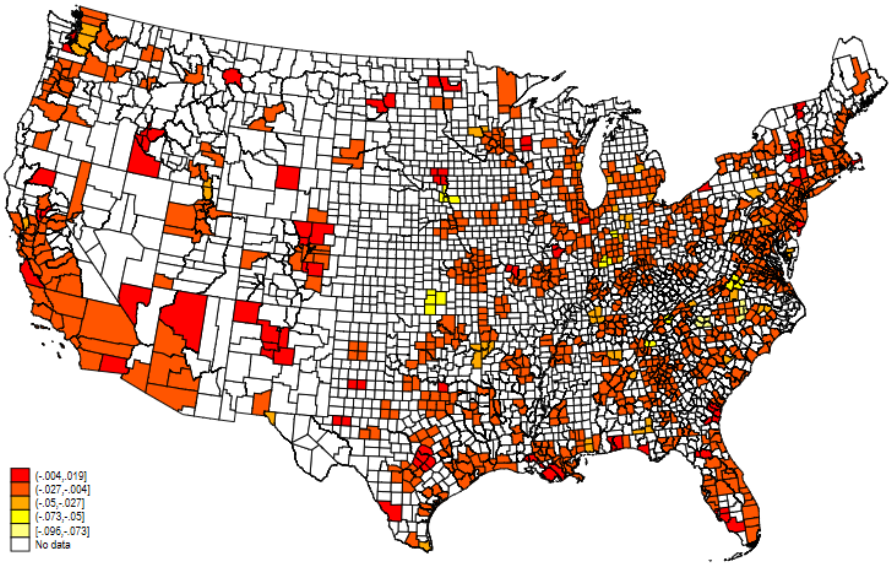
Windmeijer, Frank, 2021, Testing underidentification in linear models, with applications to dynamic panel and asset pricing models, *Journal of Econometrics*.

World Trade Organization, 2007, Sixty years of the multilateral trading system: Achievements and challenges, in *World Trade Report 2007* (Geneva: WTO).

10 Figures and Tables



(a) Net Export Growth (1991-1999)



(b) Net Export Growth (1999-2005)

Figure 9: **U.S. Mainland Metropolitan Heat Map of Net Export Growth: Two Periods**

This figure displays the U.S. mainland metropolitan heat map of net export growth measure in two periods: subfigure (a) is for 1991-1999, and subfigure (b) is for 1999-2005. The net export growth measure at the metropolitan level for a period is defined in equation (1). In the above figures, each small area with a boundary is a county. Counties with white color are non-metropolitan areas in the 2003 CBSA version (1085 counties). Metropolitan counties are painted with colors ranging from yellow (for low net export growth) to red (for high net export growth) in five categories.

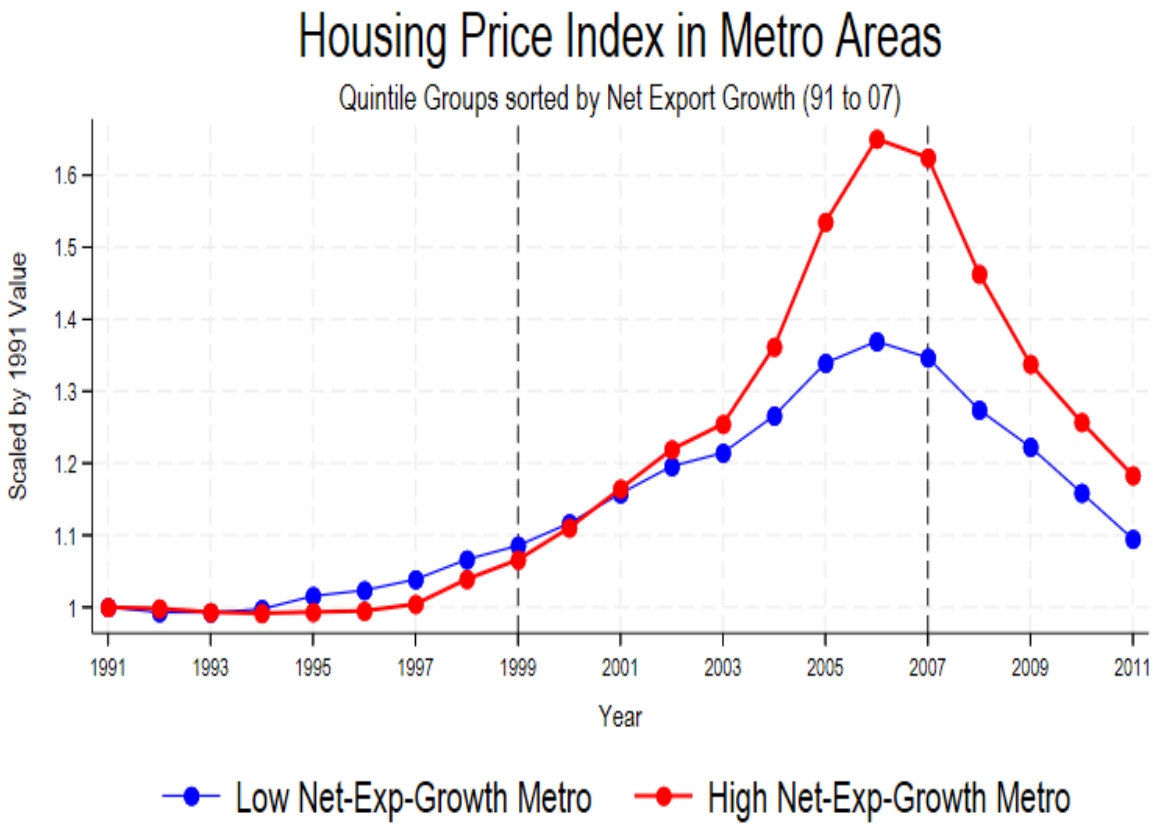


Figure 10: **Housing Price Index (91-11) in Metro Areas, High vs. Low Quintile of Net Export Growth (91 to 07)**

This figure displays the time series of weighted average housing price indices (deflated to 2007 value) for high and low quintile groups of Metropolitan Areas (MSAs) from 1991 to 2011. For the entire period, the quintile groups are sorted by net export growth (1991 to 2007) at the metropolitan level (CBSA code 2003 version). The sample includes 301 metropolitan areas (712 counties) that are consistently covered by the HMDA sample after 1990. These 679 out of consistent 712 counties have non-missing house price indexes from 1991 to 2011. The low quintile group comprises 60 metros (124 counties), and the high quintile group comprises 61 metros (92 counties) throughout the entire period. The housing price index is weighted by the county-level housing units within each group in each year. Throughout the entire period, the time series of weighted-average housing price indices of each group are divided by their 1991 values, ensuring that both groups start at a value of 1 in 1991. The red line represents the high quintile group, while the blue line represents the low quintile group.

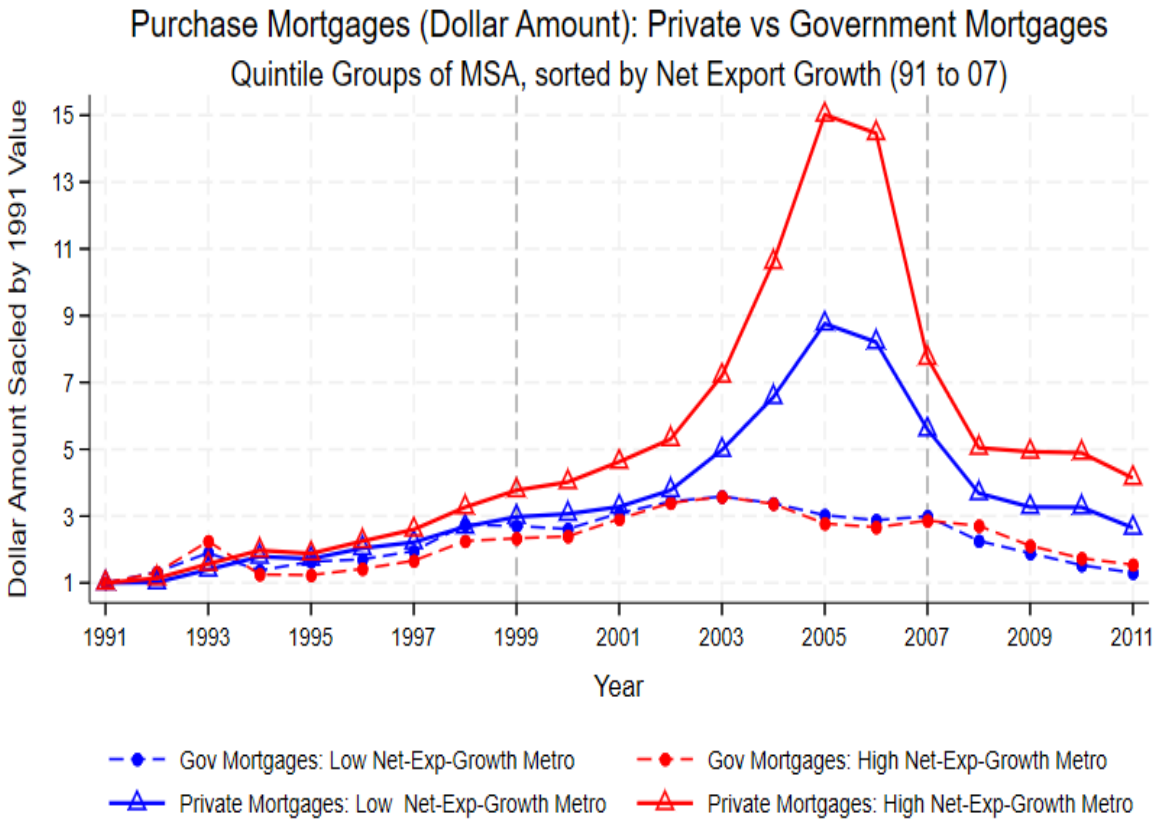


Figure 11: **Mortgage Growth (91-11) across Metropolitan Areas: GSEM vs. PLMNJ, High vs. Low Quintile Sorted by Net Export Growth (91-07).**

This figure displays the time series of the weighted-average dollar amount of Government-Sponsored Enterprise Mortgages (GSEM) (in dash lines with dots) and Private-Label Mortgages (Non-Jumbo) (PLMNJ) (in solid lines with triangles) for high and low quintile groups of metropolitan statistical areas (MSA) from 1991 to 2011. Both types of mortgages only include purchase loans. For the entire period, the quintile groups are sorted by net export growth (1991 to 2007) at the MSA level. The whole sample includes 301 MSA (712 counties) that are consistently covered by the HMDA sample after 1990 due to the smaller coverage of metropolitan areas in the early years. The low quintile group comprises 61 MSA (94 counties), and the high quintile group comprises 60 MSA (112 counties) throughout the entire period. The number of loans is weighted by the county-level housing units within each group in each year. Throughout the entire period, the time series of the weighted-average number of loans of each group are divided by their 1991 values, ensuring that both groups start at a value of 1 in 1991. The red lines represent the high quintile group, while the blue lines represent the low quintile group.

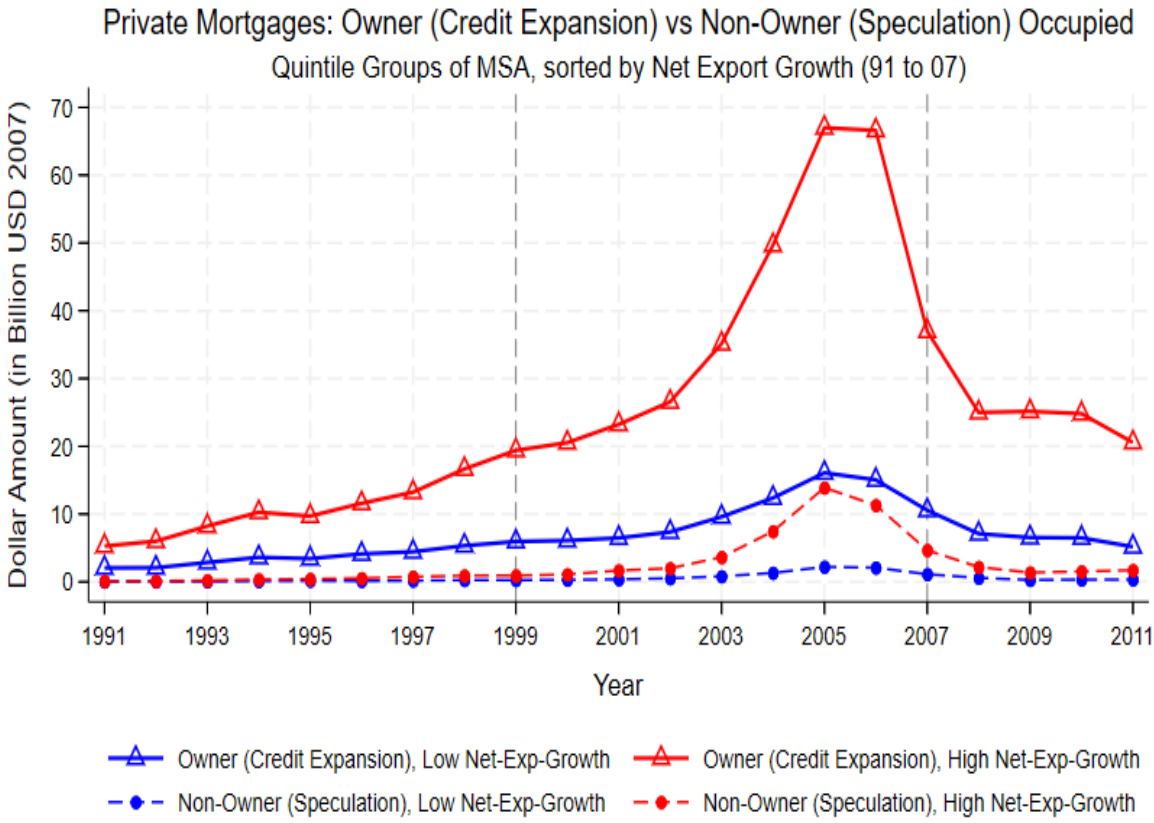
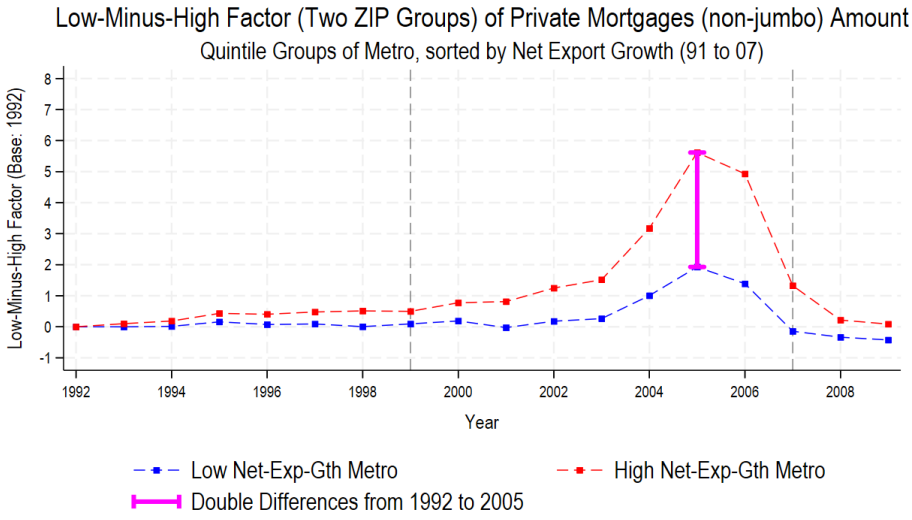
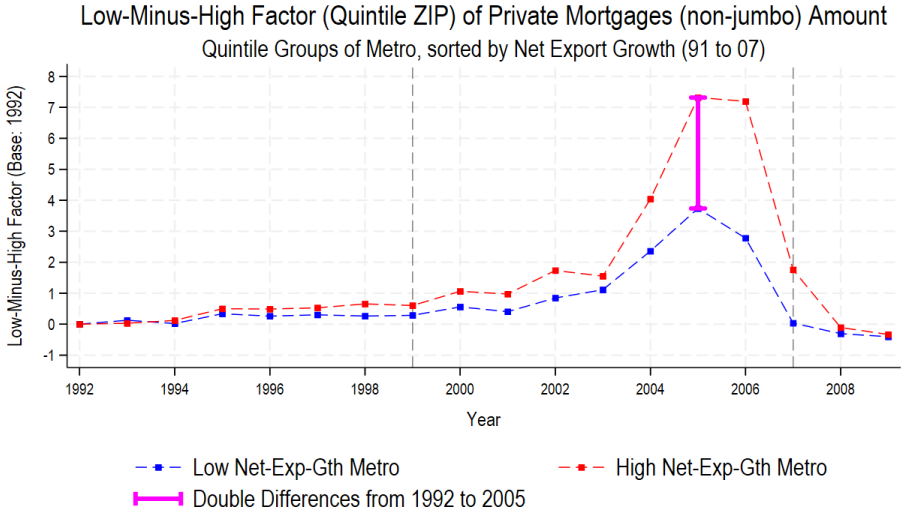


Figure 12: **Credit Expansion vs. Speculation (91-11) across Metropolitan Areas: High vs. Low Quintile Sorted by Net Export Growth (91-07).**

This figure displays the time series of dollar amount of "owner-occupied" (credit expansion) home purchase private-label mortgages (in solid lines with triangles) and "non-owner-occupied" (speculation) home purchase private-label mortgages (in dash lines with dots) from 1991 to 2011 for high and low quintile groups of metropolitan statistical areas (MSA). For the entire period, the quintile groups are sorted by net export growth (1991 to 2007) at the MSA level. The whole sample includes 301 MSA (712 counties) that are consistently covered by the HMDA sample after 1990 due to the smaller coverage of metropolitan areas in the early years. The low quintile group comprises 61 MSA (94 counties), and the high quintile group comprises 60 MSA (112 counties) throughout the entire period. The dollar amount is adjusted to the 2007 USD by the Personal Consumption Expenditures Chain-type Price Index (PCEPI) from Federal Reserve Bank of St. Louis. The red lines represent the high quintile group, while the blue lines represent the low quintile group.



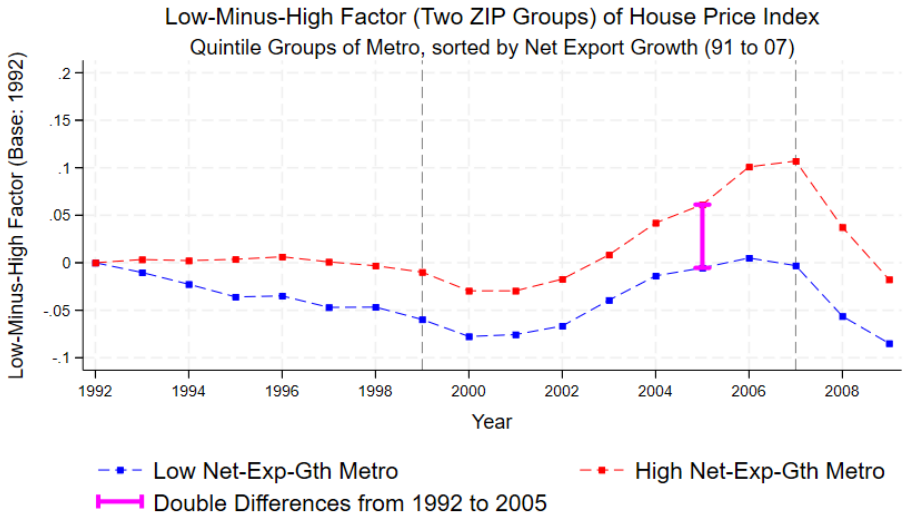
(a) Income Half Groups of ZIP Codes within Metro



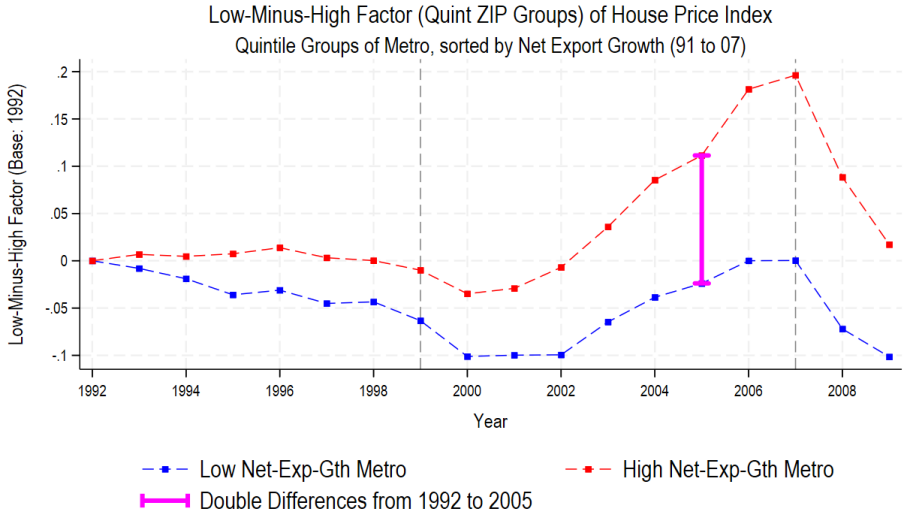
(b) Income Quintile Groups of ZIP Codes within Metro

Figure 13: **PLMNJ Amount Growth (92-09) Across ZIP Codes Across Metropolitan Areas.**

This figure displays the time series of the low-minus-high difference of the dollar amount of private-label mortgages (non-jumbo) (PLMNJ) for high (in red lines) and low (in blue lines) quintile groups of metropolitan statistical areas (MSA) from 1992 to 2009. Within each group of MSAs, the low-minus-high difference is calculated by using the dollar amount of private-label mortgage (non-jumbo) in the bottom-half/quintile-income ZIP codes to minus the top-half/quintile-income ZIP codes based on 1998 income level. Mortgages only contain home purchase loans. For the entire period, the quintile groups of MSAs are sorted by net export growth (1991 to 2007) at the MSA level. The dollar amounts are converted into 2007 US dollars. Throughout the entire period, time series mortgage amounts are divided by their 1992 values, ensuring that low-minus-high differences start at a value of 0 in 1992. For the figure of half (quintile) ZIP groups, I require that each MSA have at least three (five) ZIP codes. The whole sample for the figure of half (quintile) ZIP groups is a balanced panel of 9,969 (9,917) ZIP codes and 18 years that are consistently covered by the HMDA sample after 1990. For metropolitan areas sorted into quintile groups, the low quintile group comprises 856 (845) zip codes from 59 (56) MSAs, and the high quintile group comprises 1,527 (1,506) zip codes from 59 (53) MSAs throughout the entire period.



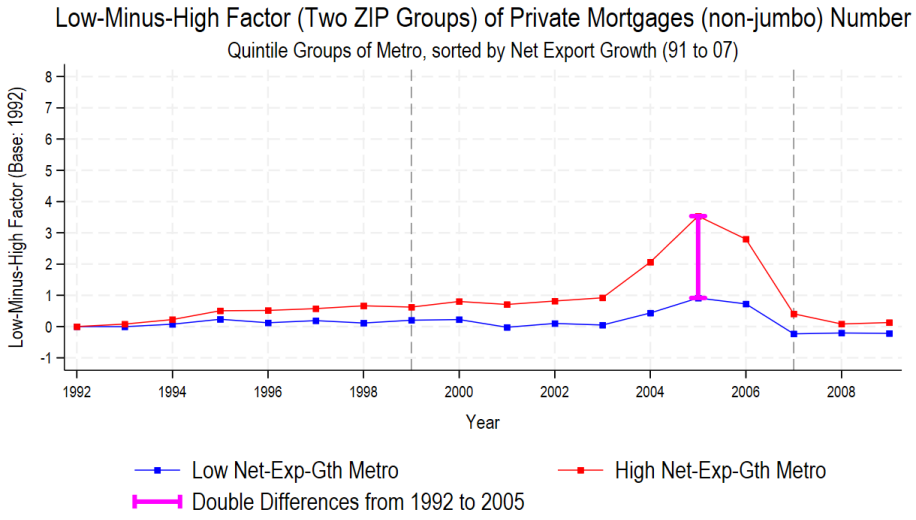
(a) Half Groups of ZIP Codes within Metro



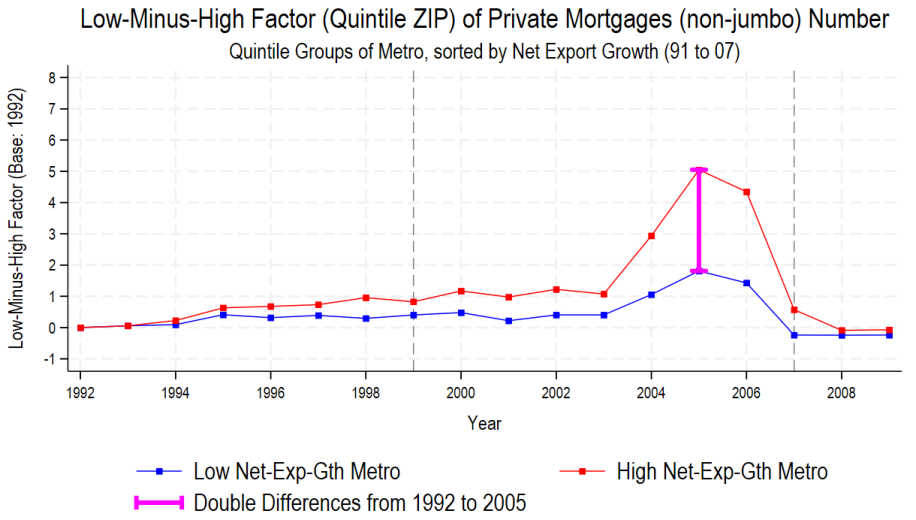
(b) Quintile Groups of ZIP Codes within Metro

Figure 14: **House Price Growth (92-09) Across ZIP Codes Across Metropolitan Areas**

This figure displays the time series of the low-minus-high difference of the weighted-average house price index for high (in red lines) and low (in blue lines) quintile groups of metropolitan statistical areas (MSAs) from 1992 to 2009. Within each group of MSAs, the low-minus-high difference is calculated by using the weighted-average house price index in the bottom-half/quintile-income ZIP codes to minus the top-half/quintile-income ZIP codes based on 1998 income level. In summing up house price index, weights are the number of house units at the zip-year level. For the entire period, the quintile groups of MSAs are sorted by net export growth (1991 to 2007) at the MSA level. Throughout the entire period, time series house price indexes are divided by their 1992 values, ensuring that low-minus-high differences start at a value of 0 in 1992. For the figure of half (quintile) ZIP groups, I require each MSA has at least three (five) ZIP codes. The whole sample for the figure of half (quintile) ZIP groups is a balanced panel of 6,083 (5,942) ZIP codes and 18 years. Compared to figures of mortgage, the reduced number of ZIP codes are due to that availability of house price index and number of house units. For metropolitan areas sorted into quintile groups, the low quintile group comprises 451 (416) zip codes from 40 (30) MSAs, and the high quintile group comprises 883 (848) zip codes from 46 (36) MSAs throughout the entire period.



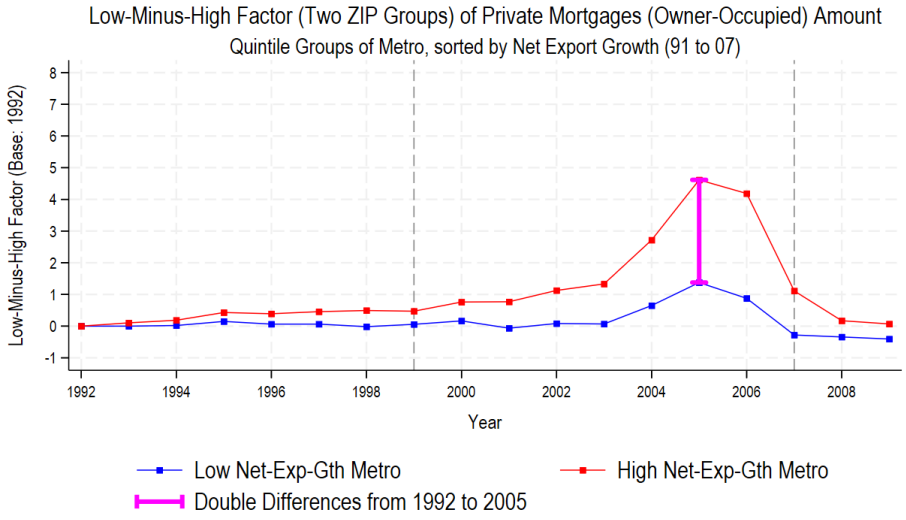
(a) Income Half Groups of ZIP Codes within Metro



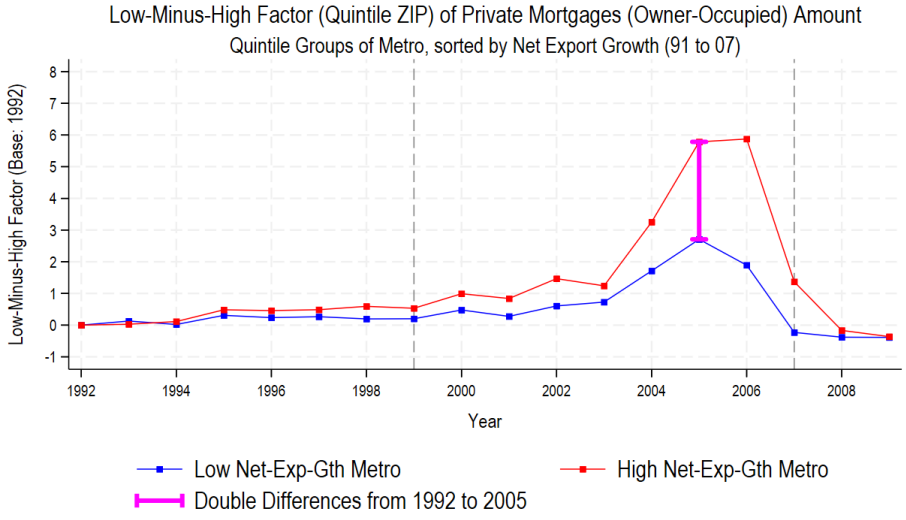
(b) Income Quintile Groups of ZIP Codes within Metro

Figure 15: **PLMNJ Number Growth (92-09) Across ZIP Codes Across Metropolitan Areas.**

This figure displays the time series of the low-minus-high difference of number of private-label mortgages (non-jumbo) (PLMNJ) for high (in red lines) and low (in blue lines) quintile groups of metropolitan statistical areas (MSA) from 1992 to 2009. Within two groups of MSAs, the low-minus-high difference is calculated by using the number of loans of private-label mortgage (non-jumbo) in the bottom-half/quintile-income ZIP codes to minus the top-half/quintile-income ZIP codes based on 1998 income level. Mortgages only contain home purchase loans. For the entire period, the quintile groups of MSAs are sorted by net export growth (1991 to 2007) at the MSA level. Throughout the entire period, time series mortgage numbers are divided by their 1992 values, ensuring that low-minus-high differences start at a value of 0 in 1992. For the figure of half (quintile) ZIP groups, I require each MSA has at least three (five) ZIP codes. The whole sample for the figure of half (quintile) ZIP groups is a balanced panel of 9,969 (9,917) ZIP codes and 18 years that are consistently covered by the HMDA sample after 1990. For metropolitan areas sorted into quintile groups, the low quintile group comprises 856 (845) zip codes from 59 (56) MSAs, and the high quintile group comprises 1,527 (1,506) zip codes from 59 (53) MSAs throughout the entire period.



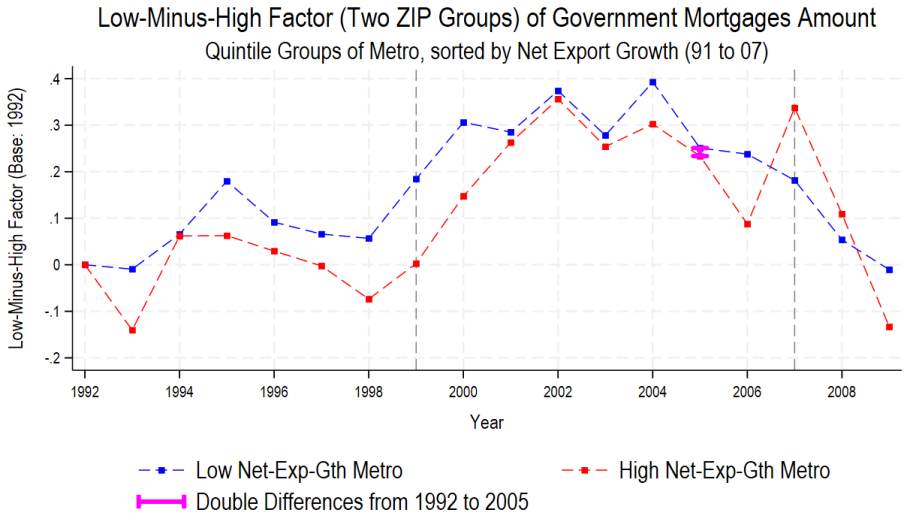
(a) Income Half Groups of ZIP Codes within Metro



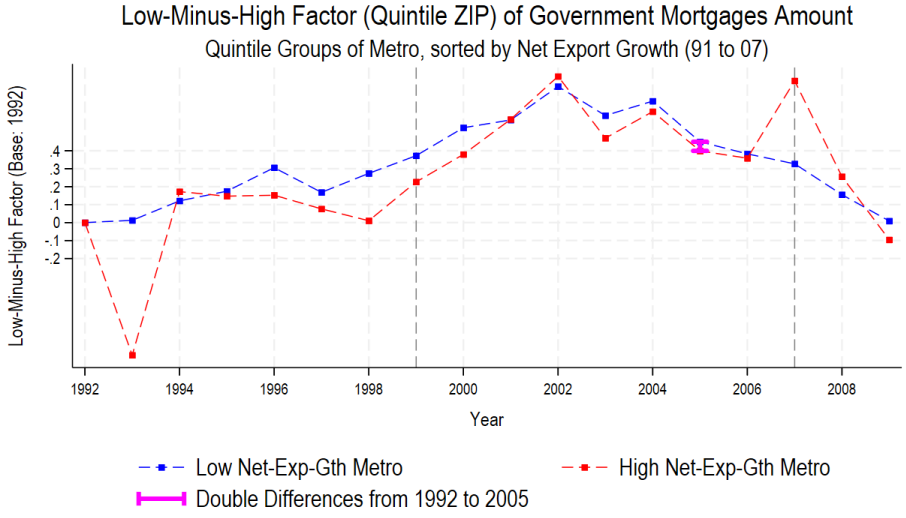
(b) Income Quintile Groups of ZIP Codes within Metro

Figure 16: **PLMNJ (Owner-Occupied) Amount Growth (92-09) Across ZIP Codes Across Metropolitan Areas.**

This figure displays the time series of the low-minus-high difference of the dollar amount of owner-occupied private-label mortgages (non-jumbo) (PLMNJ) for high (in red lines) and low (in blue lines) quintile groups of metropolitan statistical areas (MSA) from 1992 to 2009. Within each group of MSAs, the low-minus-high difference is calculated by using the dollar amount of owner-occupied private-label mortgage (non-jumbo) in the bottom-half/quintile-income ZIP codes to minus the top-half/quintile-income ZIP codes based on 1998 income level. Mortgages only contain home purchase loans. For the entire period, the quintile groups of MSAs are sorted by net export growth (1991 to 2007) at the MSA level. The dollar amounts are converted into 2007 US dollars. Throughout the entire period, time series mortgage amounts are divided by their 1992 values, ensuring that low-minus-high differences start at a value of 0 in 1992. For the figure of half (quintile) ZIP groups, I require that each MSA have at least three (five) ZIP codes. The whole sample for the figure of half (quintile) ZIP groups is a balanced panel of 9,969 (9,917) ZIP codes and 18 years that are consistently covered by the HMDA sample after 1990. For metropolitan areas sorted into quintile groups, the low quintile group comprises 856 (845) zip codes from 59 (56) MSAs, and the high quintile group comprises 1,527 (1,506) zip codes from 59 (53) MSAs throughout the entire period.



(a) Income Half Groups of ZIP Codes within Metro



(b) Income Quintile Groups of ZIP Codes within Metro

Figure 17: **GSEM Amount Growth (92-09) Across ZIP Codes Across Metropolitan Areas.**

This figure displays the time series of the low-minus-high difference of the dollar amount of government-sponsored enterprise mortgages for high (in red lines) and low (in blue lines) quintile groups of metropolitan statistical areas (MSA) from 1992 to 2009. Within each group of MSAs, the low-minus-high difference is calculated by using the dollar amount of government-sponsored enterprise mortgages in the bottom-half/quintile-income ZIP codes to minus the top-half/quintile-income ZIP codes based on 1998 income level. Mortgages only contain home purchase loans. For the entire period, the quintile groups of MSAs are sorted by net export growth (1991 to 2007) at the MSA level. The dollar amounts are converted into 2007 US dollars. Throughout the entire period, time series mortgage amounts are divided by their 1992 values, ensuring that low-minus-high differences start at a value of 0 in 1992. For the figure of half (quintile) ZIP groups, I require that each MSA have at least three (five) ZIP codes. The whole sample for the figure of half (quintile) ZIP groups is a balanced panel of 9,969 (9,917) ZIP codes and 18 years that are consistently covered by the HMDA sample after 1990. For metropolitan areas sorted into quintile groups, the low quintile group comprises 856 (845) zip codes from 59 (56) MSAs, and the high quintile group comprises 1,527 (1,506) zip codes from 59 (53) MSAs throughout the entire period.

Table 1: **Summary Statistics**

This table reports summary statistics of variables used in regressions, separated into different periods.

Variable Names	Count	Mean	SD	P25	P50	P75
Panel A. Prier Period (1991 - 1999)						
House Price Growth (07USD, 91-99, An)	672	0.0130	0.016	0.0029	0.0146	0.0236
PLMNJ Growth (07USD, 91-99, An)	705	0.1707	0.0845	0.1174	0.1638	0.2243
GSEM Growth (07USD, 91-99, An)	705	0.1495	0.0876	0.0910	0.1397	0.1904
Net Export Growth (91-99, An)	705	-0.0018	0.0017	-0.0021	-0.0013	-0.0009
GIV Net Export Growth (91-99, An)	705	-0.0004	0.0010	-0.0006	-0.0003	-0.0001
Ln(Num of households, 91)	705	10.8667	1.1624	10.0189	10.8116	11.5926
Ln(household Income, 91)	705	10.3654	0.1841	10.2467	10.3477	10.4625
Ratio of Labor Force (1989)	705	0.6694	0.0558	0.6357	0.6712	0.7053
Ln(Num of House Units, 91)	705	10.8944	1.1451	10.0776	10.8276	11.6244
Housing supply elasticity	627	2.3940	1.2058	1.5618	2.2594	3.0016
House Vacancy Rate (1989)	705	0.0827	0.0465	0.0549	0.0711	0.0961
Ratio of Renters (1989)	705	0.3158	0.0977	0.2482	0.3061	0.3690
Ratio of Bachelor Educated (1989)	705	0.1907	0.0821	0.1298	0.1754	0.2365
Ratio of White Race (1989)	705	0.8586	0.1281	0.7841	0.8983	0.9591
Ratio of Immigration (80-90)	705	0.0165	0.0248	0.0030	0.0076	0.0172
Ratio of Art, Enter, and Recre Emp (1989)	705	0.0059	0.0046	0.0041	0.0053	0.0067
Ratio of Health Emp (1989)	705	0.0392	0.0104	0.0328	0.0382	0.0448
Ratio of College Students (1989)	705	0.0097	0.0212	0.0000	0.0027	0.0105
Panel B. Boom Period (1999 - 2005)						
House Price Growth (07USD, 99-05, An)	786	0.0382	0.0342	0.0134	0.0234	0.0617
PLMNJ Growth (07USD, 99-05, An)	792	0.1698	0.0814	0.1160	0.1649	0.2164
GSEM Growth (07USD, 99-05, An)	792	0.0438	0.0711	0.0016	0.0389	0.0881
Net Export Growth (99-05, An)	792	-0.0022	0.0020	-0.0026	-0.0018	-0.0012
GIV Net Export Growth (99-05, An)	792	-0.0010	0.0011	-0.0013	-0.0008	-0.0004
Ln(Num of households, 99)	792	10.8657	1.1857	10.0668	10.8149	11.6139
Ln(household Income, 99)	792	10.7197	0.2369	10.5625	10.6946	10.8402
Fraction of Labor Force (1999)	792	0.6536	0.0552	0.6218	0.6563	0.6918
Ln(Num of House Units, 99)	792	10.9005	1.1543	10.0936	10.8644	11.6262
Housing supply elasticity	701	2.4194	1.2167	1.6055	2.2897	3.0029
House Vacancy Rate (1999)	792	0.0791	0.0494	0.0521	0.0687	0.0919
Fraction of Renters (1999)	792	0.2924	0.0977	0.2222	0.2795	0.3464
Fraction of Bachelor Educated (1999)	792	0.2058	0.0850	0.1422	0.1905	0.2517
Fraction of White Race (1999)	792	0.8215	0.1433	0.7414	0.8604	0.9357
Fraction of Immigration (90-00)	792	0.0255	0.0291	0.0066	0.0148	0.0325
Ratio of Art, Enter, and Recre Emp (1999)	792	0.0076	0.0060	0.0049	0.0065	0.0085
Ratio of Health Emp (1999)	792	0.0530	0.0114	0.0451	0.0519	0.0602
Ratio of Tradable Service Emp (1999)	792	0.0591	0.0274	0.0396	0.0524	0.0719
Ratio of College Students (1999)	792	0.0092	0.0201	0.0000	0.0024	0.0109
Panel C . Bust Period (2007 - 2009)						
House Price Growth (07USD, 07-09, An)	786	-0.0537	0.0615	-0.0701	-0.0332	-0.0158

Table 2: **2SLS Regression of House Price Growth on PLMNJ Growth in Boom Period (99-05)**

This table reports the first-stage and the second-stage results of 2SLS regression $\triangle_{99,05} Ln(HPI_c) = \beta * \triangle_{99,05} Ln(PLMNJ_c) + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05} Ln(HPI_c)$ is the growth rate of the house price index at county c 99-05, and the key variable of interest $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgage (non-jumbo) (PLMNJ) at county c 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05} Ln(PLMNJ_c)$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	PLMNJ Growth (99-05, An)	Housing Price Growth (07USD, 99-05, An)				
	(1)	(2)	(3)	(4)	(5)	(6)
GIV Net Export Growth (99-05, An)	11.361***					
	(3.392)					
PLMNJ Growth (07USD, 99-05, An)		0.500***	0.438***	0.390***	0.432***	0.438***
		(0.105)	(0.113)	(0.132)	(0.151)	(0.167)
Ln(Num of households, 99)	-0.030		0.004*	0.031**	0.027*	0.025
	(0.033)		(0.002)	(0.016)	(0.016)	(0.016)
Ln(household Income, 99)	0.012		0.054***	0.058***	0.033**	0.035**
	(0.033)		(0.012)	(0.011)	(0.015)	(0.015)
Ratio of Labor Force (1999)	0.127		-0.163***	-0.174***	-0.171***	-0.165***
	(0.103)		(0.042)	(0.050)	(0.053)	(0.051)
Ln(Num of House Units, 99)	0.041			-0.035**	-0.032**	-0.031*
	(0.033)			(0.016)	(0.016)	(0.017)
Housing supply elasticity	-0.012***			-0.005**	-0.004	-0.005*
	(0.003)			(0.002)	(0.003)	(0.003)
House Vacancy Rate (1999)	0.089			0.093**	0.072	0.092**
	(0.081)			(0.044)	(0.049)	(0.043)
Ratio of Renters (1999)	-0.112*			0.130***	0.084***	0.074**
	(0.060)			(0.024)	(0.028)	(0.038)
Ratio of Bachelor Educated (1999)	-0.104				0.066	0.024
	(0.107)				(0.046)	(0.050)
Ratio of White Race (1999)	-0.000				-0.001	-0.007
	(0.032)				(0.014)	(0.014)
Ratio of Immigration (90-00)	0.193				0.218**	0.275***
	(0.191)				(0.095)	(0.100)
Ratio of Art, Enter, and Recre Emp (1999)	1.835***					-0.291
	(0.627)					(0.397)
Ratio of Health Emp (1999)	0.172					0.363*
	(0.390)					(0.213)
Ratio of Tradable Service Emp (1999)	-0.336					0.106
	(0.319)					(0.147)
Ratio of College Students (1999)	0.021					-0.019
	(0.149)					(0.063)
Constant	-0.079	-0.046***	-0.546***	-0.532***	-0.256*	-0.286*
	(0.348)	(0.018)	(0.127)	(0.107)	(0.153)	(0.153)
Obs	695	786	786	695	695	695
R2-adj	0.127	-0.378	-0.0497	0.143	0.0363	0.0285
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
KP F-Stat	11.22	23.08	20.62	14.24	11.79	11.22
MOP F-Stat	11.20	22.02	19.84	14.40	11.90	11.20

Table 3: **Four Regressions of House Price Growth on PLMNJ Growth in Boom Period (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{99,05} Ln(HPI_c) = \beta * \triangle_{99,05} Ln(PLMNJ_c) + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05} Ln(HPI_c)$ is the growth rate of the house price index at county c 99-05, and the key variable of interest $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgage (non-jumbo) (PLMNJ) at county c 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05} Ln(PLMNJ_c)$. For the first-stage F-test, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	House Price Growth (1999-2005, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
PLMNJ Growth (07USD, 99-05, An)	0.178***	0.164***	0.123***	0.121***	0.117***
	(0.025)	(0.024)	(0.018)	(0.017)	(0.017)
R2-adj	0.167	0.345	0.508	0.528	0.539
Panel B. Reduced-form estimates					
GIV Net Export Growth (99-05, An)	7.639***	6.530***	4.655***	5.119***	4.971***
	(1.783)	(1.853)	(1.070)	(1.133)	(1.267)
R2-adj	0.0516	0.243	0.453	0.479	0.496
Panel C . 2SLS estimates					
PLMNJ Growth (07USD, 99-05, An)	0.500***	0.438***	0.390***	0.432***	0.438***
	(0.105)	(0.113)	(0.132)	(0.151)	(0.167)
Dep Var (Panel D):	PLMNJ Growth (99-05, An)				
Panel D . First-stage estimates					
GIV Net Export Growth (99-05, An)	15.291***	14.913***	11.931***	11.842***	11.361***
	(3.183)	(3.284)	(3.161)	(3.448)	(3.392)
KP F-Stat	23.08	20.62	14.24	11.79	11.22
MOP F-Stat	22.02	19.84	14.40	11.90	11.20
Controls (for all Panels)					
Basic Controls		Y	Y	Y	Y
Housing Controls			Y	Y	Y
Demographic Controls				Y	Y
Industry Controls					Y
Obs	786	786	786	786	695
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)

Table 4: **2SLS Stacked Regression of Housing Price Growth in Boom (99-05) and Bust (07-09) Periods on PLMNJ Growth (99-05)**

This table reports 2SLS regression $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = \beta_{99,05} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{07,09} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{07,09} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c)$ is the stacked growth rate of the house price index at county c 99-05 and 07-09. The key variable of interest $\triangle_{99,05}Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgages (non-jumbo) at county c 99-05. $Controls_c$ indicates control variables at county c in the period start year 1999. We use the gravity model-based instrumental variable $\triangle_{99,05}\text{givNetExp}_m$ as the IV for $\triangle_{99,05}Ln(PLMNJ_c)$. Regression is weighted by the natural logarithm of housing units in 1999. For the first-stage F-test of two non-stacked samples, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS estimates	Housing Price Growth (99-05 and 07-09, Annualized)									
	(1)		(2)		(3)		(4)		(5)	
PLMNJ Growth (07USD, 99-05, An) x Dum99t05	0.500***	(0.105)	0.438***	(0.113)	0.390***	(0.132)	0.432***	(0.151)	0.438***	(0.167)
PLMNJ Growth (07USD, 99-05, An) x Dum07t09	-0.901***	(0.209)	-0.830***	(0.211)	-0.804***	(0.273)	-0.844***	(0.261)	-0.896***	(0.293)
Dum99t05	-0.046***	(0.018)	-0.546***	(0.127)	-0.532***	(0.107)	-0.256*	(0.153)	-0.286*	(0.153)
Dum07t09	0.099***	(0.035)	0.594***	(0.171)	0.440**	(0.208)	0.726**	(0.308)	0.778**	(0.328)
Ln(Num of HH, 99) x Dum99t05			0.004*	(0.002)	0.031**	(0.016)	0.027*	(0.016)	0.025	(0.016)
Ln(Num of HH, 99) x Dum07t09			-0.009***	(0.003)	-0.070**	(0.028)	-0.027	(0.027)	-0.026	(0.029)
Ln(HH Income, 99) x Dum99t05			0.054***	(0.012)	0.058***	(0.011)	0.033**	(0.015)	0.035**	(0.015)
Ln(HH Income, 99) x Dum07t09			-0.055***	(0.017)	-0.047**	(0.021)	-0.076**	(0.030)	-0.080**	(0.031)
Ratio of Labor Force (1999) x Dum99t05			-0.163***	(0.042)	-0.174***	(0.050)	-0.171***	(0.053)	-0.165***	(0.051)
Ratio of Labor Force (1999) x Dum07t09			0.269***	(0.055)	0.341***	(0.071)	0.251***	(0.072)	0.239***	(0.080)
Ln(Num of HU, 99) x Dum99t05					-0.035**	(0.016)	-0.032**	(0.016)	-0.031*	(0.017)
Ln(Num of HU, 99) x Dum07t09					0.064**	(0.029)	0.024	(0.028)	0.023	(0.031)
Housing supply elasticity x Dum99t05					-0.005**	(0.002)	-0.004	(0.003)	-0.005*	(0.003)
Housing supply elasticity x Dum07t09					0.002	(0.004)	0.001	(0.004)	-0.000	(0.004)
House Vacancy Rate (1999) x Dum99t05					0.093**	(0.044)	0.072	(0.049)	0.092**	(0.043)
House Vacancy Rate (1999) x Dum07t09					-0.017	(0.083)	0.021	(0.081)	0.055	(0.077)
Ratio of Renters (1999) x Dum99t05					0.130***	(0.024)	0.084***	(0.028)	0.074**	(0.038)
Ratio of Renters (1999) x Dum07t09					-0.041	(0.052)	0.017	(0.048)	-0.001	(0.061)
Ratio of Bachelor Educated (1999) x Dum99t05							0.066	(0.046)	0.024	(0.050)
Ratio of Bachelor Educated (1999) x Dum07t09							0.165*	(0.090)	0.067	(0.110)
Ratio of White Race (1999) x Dum99t05							-0.001	(0.014)	-0.007	(0.014)
Ratio of White Race (1999) x Dum07t09							0.022	(0.029)	0.015	(0.030)
Ratio of Immigration (90-00) x Dum99t05							0.218**	(0.095)	0.275***	(0.100)
Ratio of Immigration (90-00) x Dum07t09							-0.562**	(0.219)	-0.527**	(0.215)
Ratio of Art, Enter, and Recre Emp (1999) x Dum99t05									-0.291	(0.397)
Ratio of Art, Enter, and Recre Emp (1999) x Dum07t09									-0.010	(0.766)
Ratio of Health Emp (1999) x Dum99t05									0.363*	(0.213)
Ratio of Health Emp (1999) x Dum07t09									0.332	(0.307)
Ratio of Tradable Service Emp (1999) x Dum99t05									0.106	(0.147)
Ratio of Tradable Service Emp (1999) x Dum07t09									0.372	(0.293)
Ratio of College Students (1999) x Dum99t05									-0.019	(0.063)
Ratio of College Students (1999) x Dum07t09									0.041	(0.130)
Obs	1572		1572		1390		1390		1390	
Cluster SE	CBSA		CBSA		CBSA		CBSA		CBSA	
Weight	Ln(HU99)		Ln(HU99)		Ln(HU99)		Ln(HU99)		Ln(HU99)	
KP F-Stat (99-05, non-stack sample)	23.08		20.62		14.24		11.79		11.22	
MOP F-Stat (99-05, non-stack sample)	22.02		19.84		14.40		11.90		11.20	
CoefEqual_Chi2	23.100		17.863		9.835		10.716		9.664	
CoefEqual_PValue	0.000		0.000		0.002		0.001		0.002	

Table 5: **Four Stacked Regressions of Housing Price Growth in Boom (99-05) and Bust (07-09) Periods on PLMNJ Growth (99-05)**

This table reports OLS, reduced-form, first stage, and second stages of stacked 2SLS regression $\triangle_{99,05}$ & $\triangle_{07,09}Ln(HPI_c) = \beta_{99,05} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{07,09} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{07,09} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c)$ is the stacked growth rate of the house price index at county c 99-05 and 07-09. The key variable of interest $\triangle_{99,05}Ln(PLMNJ_c$ is the growth rate of the dollar amount of private-label mortgages (non-jumbo) at county c 99-05. $Controls_c$ indicates control variables at county c in the period start year 1999. We use the gravity model-based instrumental variable $\triangle_{99,05}\text{givNetExp}_m$ as the IV for $\triangle_{99,05}Ln(PLMNJ_c$. Each regression is weighted by the natural logarithm of housing units in the start year (either 1999 or 2007). For the first-stage F-test of the non-stacked sample, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	House Price Growth (99-05 & 07-09, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
PLMNJ Growth (07USD, 99-05, An) x Dum99t05	0.178***	0.164***	0.123***	0.121***	0.117***
	(0.025)	(0.024)	(0.018)	(0.017)	(0.017)
PLMNJ Growth (07USD, 99-05, An) x Dum07t09	-0.399***	-0.375***	-0.303***	-0.287***	-0.267***
	(0.051)	(0.050)	(0.046)	(0.041)	(0.040)
R2-adj	0.595	0.666	0.703	0.735	0.746
Panel B. Reduced-form estimates					
GIV Net Export Growth (99-05, An) x Dum99t05	7.639***	6.530***	4.655***	5.119***	4.971***
	(1.784)	(1.853)	(1.070)	(1.133)	(1.267)
GIV Net Export Growth (99-05, An) x Dum07t09	-13.773***	-12.383***	-9.595***	-10.000***	-10.174***
	(2.768)	(2.839)	(2.151)	(2.092)	(2.083)
R2-adj	0.499	0.579	0.652	0.691	0.711
Panel C . 2SLS estimates					
PLMNJ Growth (07USD, 99-05, An) x Dum99t05	0.500***	0.438***	0.390***	0.432***	0.438***
	(0.105)	(0.113)	(0.132)	(0.151)	(0.167)
PLMNJ Growth (07USD, 99-05, An) x Dum07t09	-0.901***	-0.830***	-0.804***	-0.844***	-0.896***
	(0.209)	(0.211)	(0.273)	(0.261)	(0.293)
Dep Var (Panel D):	PLMNJ Growth (99-05, An)				
Panel D . First-stage estimates only for 99-05 (Non-stack sample)					
GIV Net Export Growth (99-05, An)	16.743***	15.759***	15.431***	15.436***	14.262***
	(3.177)	(3.177)	(3.097)	(3.094)	(3.059)
KP F-Stat	23.08	20.62	14.24	11.79	11.22
MOP F-Stat	22.02	19.84	14.40	11.90	11.20
Controls (for all Panels)					
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs (Panel A, B, and C)	1572	1572	1390	1390	1390
Obs (Panel D)	786	786	695	695	695
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)

Table 6: **2SLS Regression of House Price Growth on PLMNJ Growth with Two Instruments in Boom Period (99-05)**

This table reports the first-stage and the second-stage results of 2SLS regression $\triangle_{99,05} Ln(HPI_c) = \beta * \triangle_{99,05} Ln(PLMNJ_c) + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05} Ln(HPI_c)$ is the growth rate of the house price index at county c 99-05, and the key variable of interest $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgage (non-jumbo) (PLMNJ) at county c 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) and its square as two IVs for $\triangle_{99,05} Ln(PLMNJ_c)$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	PLMNJ Growth (99-05, An)	Housing Price Growth (07USD, 99-05, An)				
	(1)	(2)	(3)	(4)	(5)	(6)
GIV Net Export Growth (99-05, An)	9.492					
	(7.761)					
GIV Net Export Growth Suqare (99-05, An)	-356.147					
	(1311.626)					
PLMNJ Growth (07USD, 99-05, An)		0.447***	0.411***	0.389***	0.444***	0.437***
		(0.099)	(0.114)	(0.131)	(0.150)	(0.163)
Ln(Num of households, 99)	-0.030		0.004**	0.031**	0.026	0.024
	(0.033)		(0.002)	(0.016)	(0.016)	(0.016)
Ln(household Income, 99)	0.012		0.054***	0.059***	0.034**	0.036**
	(0.033)		(0.013)	(0.011)	(0.015)	(0.015)
Ratio of Labor Force (1999)	0.127		-0.172***	-0.177***	-0.175***	-0.169***
	(0.103)		(0.043)	(0.049)	(0.052)	(0.050)
Ln(Num of House Units, 99)	0.040			-0.034**	-0.031*	-0.030*
	(0.033)			(0.015)	(0.016)	(0.016)
Housing supply elasticity	-0.012***			-0.005**	-0.004	-0.005*
	(0.003)			(0.002)	(0.003)	(0.003)
House Vacancy Rate (1999)	0.090			0.089**	0.067	0.087**
	(0.081)			(0.043)	(0.048)	(0.041)
Ratio of Renters (1999)	-0.110*			0.130***	0.085***	0.077**
	(0.060)			(0.023)	(0.028)	(0.037)
Ratio of Bachelor Educated (1999)	-0.101				0.067	0.023
	(0.108)				(0.046)	(0.049)
Ratio of White Race (1999)	-0.000				-0.001	-0.006
	(0.032)				(0.014)	(0.014)
Ratio of Immigration (90-00)	0.193				0.211**	0.258***
	(0.193)				(0.094)	(0.096)
Ratio of Art, Enter, and Recre Emp (1999)	1.845***					-0.275
	(0.632)					(0.388)
Ratio of Health Emp (1999)	0.171					0.332
	(0.391)					(0.206)
Ratio of Tradable Service Emp (1999)	-0.345					0.112
	(0.320)					(0.143)
Ratio of College Students (1999)	0.022					-0.022
	(0.150)					(0.062)
Constant	-0.079	-0.036**	-0.547***	-0.541***	-0.268*	-0.294**
	(0.348)	(0.016)	(0.130)	(0.105)	(0.152)	(0.149)
Obs	695	786	786	695	695	695
R2-adj	0.126	-0.215	0.0224	0.146	-0.000824	0.0283
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
KP F-Stat	5.691	13.23	12.07	7.144	5.890	5.691
MOP F-Stat	4.835	10.64	9.880	6.337	5.275	4.835
Hansen J-Test Stat		1.590	1.794	0.559	0.453	0.698
Hansen J-Test P-value		0.207	0.180	0.455	0.501	0.403

Table 7: **Four Regressions of House Price Growth on PLMNJ Growth in Boom Period (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{99,05} Ln(HPI_c) = \beta * \triangle_{99,05} Ln(PLMNJ_c) + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05} Ln(HPI_c)$ is the growth rate of the house price index at county c 99-05, and the key variable of interest $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgage (non-jumbo) (PLMNJ) at county c 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) and its square as two IVs for $\triangle_{99,05} Ln(PLMNJ_c)$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	House Price Growth (1999-2005, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
PLMNJ Growth (07USD, 99-05, An)	0.178***	0.164***	0.123***	0.121***	0.117***
	(0.025)	(0.024)	(0.018)	(0.017)	(0.017)
R2-adj	0.167	0.345	0.508	0.528	0.539
Panel B. Reduced-form estimates					
GIV Net Export Growth (99-05, An)	17.955**	18.327**	11.397	11.940	9.492
	(7.847)	(8.035)	(7.340)	(7.706)	(7.761)
GIV Net Export Growth Suqare (99-05, An)	521.917	669.230	-100.976	18.654	-356.147
	(1187.665)	(1195.293)	(1217.866)	(1311.892)	(1311.626)
R2-adj	0.0381	0.0425	0.105	0.108	0.126
Panel C . 2SLS estimates					
PLMNJ Growth (07USD, 99-05, An)	0.447***	0.411***	0.389***	0.444***	0.437***
	(0.099)	(0.114)	(0.131)	(0.150)	(0.163)
Dep Var (Panel D):	PLMNJ Growth (99-05, An)				
Panel D . First-stage estimates					
GIV Net Export Growth (99-05, An)	17.955**	18.327**	11.397	11.940	9.492
	(7.847)	(8.035)	(7.340)	(7.706)	(7.761)
GIV Net Export Growth Suqare (99-05, An)	521.917	669.230	-100.976	18.654	-356.147
	(1187.665)	(1195.293)	(1217.866)	(1311.892)	(1311.626)
KP F-Stat	13.23	12.07	7.144	5.890	5.691
MOP F-Stat	10.64	9.880	6.337	5.275	4.835
Controls (for all Panels)					
Basic Controls		Y	Y	Y	Y
Housing Controls			Y	Y	Y
Demographic Controls				Y	Y
Industry Controls					Y
Obs	786	786	786	786	695
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)

Table 8: **2SLS Regression of PLMNJ Growth on Net Export Growth in Boom Period (99-05)**

This table reports the first-stage and the second-stage results of 2SLS regression $\triangle_{99,05} Ln(PLMNJ_c) = \beta * \triangle_{99,05}\text{NetExp}_m + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgages (non-jumbo) (PLMNJ) at county c 99-05 and the key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in 1999 as the weight for the housing market. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	NEG (99-05, An)	Private-label Mortgage (non-jumbo) Growth (99-05, 07USD, An)				
	(1)	(2)	(3)	(4)	(5)	(6)
GIV Net Export Growth (99-05, An)	1.066***					
	(0.258)					
Net Export Growth (99-05, An)		13.694***	13.365***	11.145***	10.851***	10.992***
		(3.547)	(3.643)	(3.254)	(3.331)	(3.231)
Ln(Num of households, 99)	0.000		0.006**	0.021	0.003	-0.007
	(0.001)		(0.003)	(0.032)	(0.031)	(0.031)
Ln(household Income, 99)	-0.003***		-0.004	-0.027	0.035	0.043
	(0.001)		(0.022)	(0.025)	(0.036)	(0.034)
Fraction of Labor Force (1999)	-0.004		-0.021	0.081	0.127	0.165
	(0.002)		(0.107)	(0.099)	(0.102)	(0.102)
Ln(Num of House Units, 99)	-0.000			-0.010	0.009	0.021
	(0.001)			(0.033)	(0.032)	(0.032)
Housing supply elasticity	-0.000*			-0.011***	-0.011***	-0.010***
	(0.000)			(0.003)	(0.003)	(0.003)
House Vacancy Rate (1999)	0.001			0.179**	0.187**	0.106
	(0.002)			(0.087)	(0.084)	(0.084)
Fraction of Renters (1999)	-0.000			-0.107**	-0.054	-0.113*
	(0.001)			(0.053)	(0.058)	(0.058)
Fraction of Bachelor Educated (1999)	0.006**				-0.239**	-0.176
	(0.003)				(0.094)	(0.108)
Fraction of White Race (1999)	0.000				0.018	0.000
	(0.001)				(0.035)	(0.034)
Fraction of Immigration (90-00)	-0.005				0.138	0.250
	(0.004)				(0.197)	(0.200)
Ratio of Art, Enter, and Recre Emp (1999)	0.014					1.726***
	(0.009)					(0.631)
Ratio of Health Emp (1999)	-0.014					0.410
	(0.011)					(0.410)
Ratio of Tradable Service Emp (1999)	0.021***					-0.582*
	(0.007)					(0.327)
Ratio of College Students (1999)	-0.002					0.103
	(0.003)					(0.142)
Constant	0.030***	0.200***	0.181	0.352	-0.339	-0.446
	(0.009)	(0.009)	(0.212)	(0.236)	(0.371)	(0.358)
Obs	701	792	792	701	701	701
R2-adj	0.413	0.0239	0.0309	0.0881	0.100	0.124
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
KP F-Stat	17.11	23.36	23.11	16.44	17.78	17.11
MOP F-Stat	16.83	22.91	22.63	16.08	17.44	16.83

Table 9: **Four Regressions of PLMNJ Growth on Net Export Growth in Boom Period (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{99,05} Ln(PLMNJ_c) = \beta * \triangle_{99,05}\text{NetExp}_m + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05} Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgages (non-jumbo) (PLMNJ) at county c 99-05 and the key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Private-label Mortgage (non-jumbo) Growth (07USD, 99-05, An)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
Net Export Growth (99-05, An)	8.391***	8.110***	5.993***	6.564***	6.786***
	(1.551)	(1.574)	(1.627)	(1.728)	(1.671)
R2-adj	0.0406	0.0473	0.104	0.112	0.134
Panel B. Reduced-form estimates					
GIV Net Export Growth (99-05, An)	15.753***	15.278***	12.243***	12.134***	11.716***
	(3.255)	(3.331)	(3.276)	(3.541)	(3.465)
R2-adj	0.0407	0.0477	0.109	0.112	0.132
Panel C . 2SLS estimates					
Net Export Growth (99-05, An)	13.694***	13.365***	11.145***	10.851***	10.992***
	(3.547)	(3.643)	(3.254)	(3.331)	(3.231)
Dep Var (Panel D):	Net Export Growth (99-05, An)				
Panel D . First-stage estimates					
GIV Net Export Growth (99-05, An)	1.150***	1.143***	1.098***	1.118***	1.066***
	(0.238)	(0.238)	(0.271)	(0.265)	(0.258)
KP F-Stat	23.36	23.11	16.44	17.78	17.11
MOP F-Stat	22.91	22.63	16.08	17.44	16.83
Controls (for all Panels)					
Basic Controls		Y	Y	Y	Y
Housing Controls			Y	Y	Y
Demographic Controls				Y	Y
Industry Controls					Y
Obs	792	792	792	792	701
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)

Table 10: **2SLS Stacked Regression of GSEM and PLMNJ Growth in Boom Period (99-05) on Net Export Growth (99-05)**

This table reports 2SLS stacked regression $\triangle_{99,05}Ln(GSEMnPLMNJ_c) = \beta_G * \triangle_{99,05}\text{NetExp}_m \times Dum_G + \beta_P * \triangle_{99,05}\text{NetExp}_m \times Dum_P + \gamma_G * \boldsymbol{Controls_c} \times Dum_G + \gamma_P * \boldsymbol{Controls_c} \times Dum_P + \alpha_G + \alpha_P + \epsilon_{G,c} + \epsilon_{P,c}$. The left-hand-side dependent variable $\triangle_{99,05}Ln(GSEMnPLMNJ_c)$ is the stacked growth rate of the dollar amount of either government-sponsored enterprise mortgages (GSEM) or private-label mortgages (non-jumbo) (PLMNJ) at county c 99-05. The key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in 1999. We report the statistics and p-values for the tests of coefficient equality between β_G and β_P. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS estimates	GSEM and PLMNJ Growth (1999-2005, annualized)									
	(1)		(2)		(3)		(4)		(5)	
Net Export Growth (99-05, An) x Dum_GSEM	1.889	(2.539)	4.626**	(2.265)	3.071	(2.329)	3.149	(2.307)	2.205	(2.470)
Net Export Growth (99-05, An) x Dum_PLMNJ	13.694***	(3.547)	13.365***	(3.643)	11.145***	(3.254)	10.851***	(3.331)	10.992***	(3.231)
Dum_GSEM	0.046***	(0.008)	1.074***	(0.187)	0.851***	(0.207)	1.216***	(0.329)	1.289***	(0.339)
Dum_PLMNJ	0.200***	(0.009)	0.181	(0.212)	0.352	(0.236)	-0.339	(0.371)	-0.446	(0.358)
Ln(Num of HH, 99) x Dum_GSEM			-0.015***	(0.002)	0.011	(0.036)	0.054	(0.034)	0.047	(0.032)
Ln(Num of HH, 99) x Dum_PLMNJ			0.006**	(0.003)	0.021	(0.032)	0.003	(0.031)	-0.007	(0.031)
Ln(HH Income, 99) x Dum_GSEM			-0.067***	(0.020)	-0.049**	(0.023)	-0.090***	(0.033)	-0.097***	(0.033)
Ln(HH Income, 99) x Dum_PLMNJ			-0.004	(0.022)	-0.027	(0.025)	0.035	(0.036)	0.043	(0.034)
Ratio of Labor Force (1999) x Dum_GSEM			-0.214***	(0.075)	-0.228**	(0.091)	-0.327***	(0.084)	-0.334***	(0.089)
Ratio of Labor Force (1999) x Dum_PLMNJ			-0.021	(0.107)	0.081	(0.099)	0.127	(0.102)	0.165	(0.102)
Ln(Num of HU, 99) x Dum_GSEM					-0.023	(0.037)	-0.065*	(0.035)	-0.059*	(0.033)
Ln(Num of HU, 99) x Dum_PLMNJ					-0.010	(0.033)	0.009	(0.032)	0.021	(0.032)
Housing supply elasticity x Dum_GSEM					0.002	(0.003)	0.002	(0.003)	0.002	(0.003)
Housing supply elasticity x Dum_PLMNJ					-0.011***	(0.003)	-0.011***	(0.003)	-0.010***	(0.003)
House Vacancy Rate (1999) x Dum_GSEM					0.216**	(0.093)	0.247***	(0.094)	0.232**	(0.096)
House Vacancy Rate (1999) x Dum_PLMNJ					0.179**	(0.087)	0.187**	(0.084)	0.106	(0.084)
Ratio of Renters (1999) x Dum_GSEM					-0.032	(0.061)	0.045	(0.061)	-0.012	(0.065)
Ratio of Renters (1999) x Dum_PLMNJ					-0.107**	(0.053)	-0.054	(0.058)	-0.113*	(0.058)
Ratio of Bachelor Educated (1999) x Dum_GSEM							0.172**	(0.083)	0.066	(0.096)
Ratio of Bachelor Educated (1999) x Dum_PLMNJ							-0.239**	(0.094)	-0.176	(0.108)
Ratio of White Race (1999) x Dum_GSEM							0.085***	(0.024)	0.063**	(0.024)
Ratio of White Race (1999) x Dum_PLMNJ							0.018	(0.035)	0.000	(0.034)
Ratio of Immigration (90-00) x Dum_GSEM							-0.326	(0.202)	-0.230	(0.196)
Ratio of Immigration (90-00) x Dum_PLMNJ							0.138	(0.197)	0.250	(0.200)
Ratio of Art, Enter, and Recre Emp (1999) x Dum_GSEM									1.176***	(0.356)
Ratio of Art, Enter, and Recre Emp (1999) x Dum_PLMNJ									1.726***	(0.631)
Ratio of Health Emp (1999) x Dum_GSEM									0.733**	(0.294)
Ratio of Health Emp (1999) x Dum_PLMNJ									0.410	(0.410)
Ratio of Tradable Service Emp (1999) x Dum_GSEM									0.305	(0.339)
Ratio of Tradable Service Emp (1999) x Dum_PLMNJ									-0.582*	(0.327)
Ratio of College Students (1999) x Dum_GSEM									0.130	(0.133)
Ratio of College Students (1999) x Dum_PLMNJ									0.103	(0.142)
Obs	1584		1584		1402		1402		1402	
Cluster SE	CBSA		CBSA		CBSA		CBSA		CBSA	
Weight	Ln(HU99)		Ln(HU99)		Ln(HU99)		Ln(HU99)		Ln(HU99)	
KP F-Stat (Non-Stack Sample)	23.36		23.11		16.44		17.78		17.11	
MOP F-Stat (Non-Stack Sample)	22.91		22.63		16.08		17.44		16.83	
CoefEqual_Chi2	8.165		7.083		6.247		6.022		6.646	
CoefEqual_PValue	0.004		0.008		0.012		0.014		0.010	

Table 11: **Four Stacked Regressions of GSEM and PLMNJ Growth in Boom Period (99-05) on Net Export Growth (99-05)**

This table reports OLS, reduced-form, and the second stage of stacked 2SLS regression $\triangle_{99,05}Ln(GSEMnPLMNJ_c) = \beta_G * \triangle_{99,05}\text{NetExp}_m \times Dum_G + \beta_P * \triangle_{99,05}\text{NetExp}_m \times Dum_P + \gamma_G * \boldsymbol{Controls_c} \times Dum_G + \gamma_P * \boldsymbol{Controls_c} \times Dum_P + \alpha_G + \alpha_P + \epsilon_{G,c} + \epsilon_{P,c}$. The left-hand-side dependent variable $\triangle_{99,05}Ln(GSEMnPLMNJ_c)$ is the stacked growth rate of the dollar amount of either government-sponsored enterprise mortgages (GSEM) or private-label mortgages (non-jumbo) (PLMNJ) at county c 99-05. The key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var	GSEM and PLMNJ Growth (99-05, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
Net Export Growth (99-05, An) x Dum_GSEM	2.928**	4.552***	3.800***	3.286**	2.850**
	(1.360)	(1.380)	(1.368)	(1.338)	(1.308)
Net Export Growth (99-05, An) x Dum_PLMNJ	8.391***	8.110***	5.993***	6.564***	6.786***
	(1.552)	(1.575)	(1.627)	(1.729)	(1.671)
R2-adj	0.733	0.761	0.777	0.782	0.787
Panel B. Reduced-form estimates					
GIV Net Export Growth (99-05, An) x Dum_GSEM	2.173	5.288**	3.373	3.522	2.350
	(2.987)	(2.567)	(2.633)	(2.515)	(2.673)
GIV Net Export Growth (99-05, An) x Dum_PLMNJ	15.753***	15.278***	12.243***	12.134***	11.716***
	(3.256)	(3.332)	(3.277)	(3.543)	(3.466)
R2-adj	0.733	0.760	0.777	0.782	0.786
Panel C . 2SLS estimates					
Net Export Growth (99-05, An) x Dum_GSEM	1.889	4.626**	3.071	3.149	2.205
	(2.539)	(2.265)	(2.329)	(2.307)	(2.470)
Net Export Growth (99-05, An) x Dum_PLMNJ	13.694***	13.365***	11.145***	10.851***	10.992***
	(3.547)	(3.643)	(3.254)	(3.331)	(3.231)
Dep Var (Panel D):	Net Export Growth (99-05, An)				
Panel D . First-stage estimates for non-stack sample					
GIV Net Export Growth (99-05, An)	1.150***	1.143***	1.098***	1.118***	1.066***
	(0.238)	(0.238)	(0.271)	(0.265)	(0.258)
KP F-Stat	23.36	23.11	16.44	17.78	17.11
MOP F-Stat	22.91	22.63	16.08	17.44	16.83
Controls (for all Panels)					
Dum_MortgageType	Y	Y	Y	Y	Y
Basic Controls x Dum_MortgageType		Y	Y	Y	Y
Housing Controls x Dum_MortgageType			Y	Y	Y
Demographic Controls x Dum_MortgageType				Y	Y
Industry Controls x Dum_MortgageType					Y
Obs (Panel A, B, & C)	1584	1584	1402	1402	1402
Obs (Panel D)	792	792	701	701	701
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)

Table 12: **2SLS Stacked Regression of House Price Growth on Net Export Growth in Prior (91-99) and Boom (99-05) Periods**

This table reports 2SLS stacked regression $\triangle_{91,99}\&\triangle_{99,05}Ln(HPI_c) = \beta_{91,99} * \triangle_{91,99}\text{NetExp}_m \times Dum_{91,99} + \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \gamma_{91,99} * \boldsymbol{Controls_c} \times Dum_{91,99} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \alpha_{91,99} + \alpha_{99,05} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{91,99}\&\triangle_{99,05}Ln(HPI_c)$ is the stacked growth rate of house price index at county c 91-99 and 99-05. The key variable of interest $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$ are the growth rate of net export at the metropolitan area (CBSA03 code) m 91-99 and 99-05, respectively. $Controls_c$ indicates control variables at county c in the period start year either 1991 or 1999. We use the gravity model-based instrumental variable $\triangle_{91,99}\text{givNetExp}_m$ and $\triangle_{99,05}\text{givNetExp}_m$ as IVs for $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$. Regression is weighted by the natural logarithm of housing units in start year (either 1991 or 1999). We report the statistics and p-values for the tests of coefficient equality between $\beta_{91,99}$ and $\beta_{99,05}$. For the first-stage F-test of two non-stacked samples, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. We report Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS estimates	House Price Growth (07USD, 91-99 & 99-05, annualized)									
	(1)		(2)		(3)		(4)		(5)	
Net Export Growth (91-99, An) x Dum91t99	-3.530**	(1.684)	-3.125*	(1.687)	-2.037	(1.719)	-1.736	(1.743)	-1.902	(1.767)
Net Export Growth (99-05, An) x Dum99t05	6.926***	(2.327)	6.020***	(2.316)	4.470***	(1.725)	4.976***	(1.805)	4.884***	(1.780)
Dum91t99	0.007**	(0.003)	0.183***	(0.046)	0.181***	(0.049)	0.286***	(0.081)	0.323***	(0.095)
Dum99t05	0.055***	(0.006)	-0.478***	(0.081)	-0.429***	(0.079)	-0.464***	(0.131)	-0.502***	(0.128)
Ln(Num of HH, 91) x Dum91t99			-0.002***	(0.001)	-0.043***	(0.011)	-0.034***	(0.010)	-0.031***	(0.010)
Ln(Num of HH, 99) x Dum99t05			0.007***	(0.002)	0.031**	(0.014)	0.014	(0.013)	0.008	(0.013)
Ln(HH Income, 91) x Dum91t99			-0.017***	(0.005)	-0.018***	(0.005)	-0.030***	(0.008)	-0.032***	(0.009)
Ln(HH Income, 99) x Dum99t05			0.053***	(0.008)	0.052***	(0.008)	0.055***	(0.012)	0.058***	(0.012)
Ratio of Labor Force (1989) x Dum91t99			0.049**	(0.021)	0.066***	(0.023)	0.057***	(0.021)	0.030	(0.021)
Ratio of Labor Force (1999) x Dum99t05			-0.192***	(0.042)	-0.154***	(0.043)	-0.120***	(0.042)	-0.109***	(0.039)
Ln(Num of HU, 91) x Dum91t99					0.042***	(0.011)	0.033***	(0.010)	0.030***	(0.010)
Ln(Num of HU, 99) x Dum99t05					-0.031**	(0.014)	-0.015	(0.013)	-0.010	(0.013)
Housing supply elasticity x Dum91t99					0.000	(0.001)	0.000	(0.001)	0.001	(0.001)
Housing supply elasticity x Dum99t05					-0.010***	(0.002)	-0.009***	(0.002)	-0.010***	(0.002)
House Vacancy Rate (1989) x Dum91t99					-0.098***	(0.025)	-0.093***	(0.024)	-0.105***	(0.026)
House Vacancy Rate (1999) x Dum99t05					0.165***	(0.037)	0.153***	(0.038)	0.136***	(0.040)
Ratio of Renters (1989) x Dum91t99					-0.022**	(0.010)	-0.016	(0.013)	-0.001	(0.014)
Ratio of Renters (1999) x Dum99t05					0.100***	(0.018)	0.070**	(0.028)	0.033	(0.029)
Ratio of Bachelor Educated (1989) x Dum91t99							0.036**	(0.016)	0.069***	(0.024)
Ratio of Bachelor Educated (1999) x Dum99t05							-0.052	(0.035)	-0.092**	(0.036)
Ratio of White Race (1989) x Dum91t99							0.010	(0.009)	0.015	(0.009)
Ratio of White Race (1999) x Dum99t05							0.005	(0.015)	-0.011	(0.015)
Ratio of Immigration (80-90) x Dum91t99							-0.097***	(0.037)	-0.135***	(0.038)
Ratio of Immigration (90-00) x Dum99t05							0.287***	(0.077)	0.380***	(0.076)
Ratio of Art, Enter, and Recre Emp (1989) x Dum91t99									0.128	(0.149)
Ratio of Art, Enter, and Recre Emp (1999) x Dum99t05									0.526***	(0.193)
Ratio of Health Emp (1989) x Dum91t99									-0.253***	(0.088)
Ratio of Health Emp (1999) x Dum99t05									0.585***	(0.133)
Ratio of College Students (1989) x Dum91t99									-0.121*	(0.064)
Ratio of College Students (1999) x Dum99t05									0.010	(0.053)
Obs	1344		1344		1198		1198		1198	
Cluster SE	CBSA		CBSA		CBSA		CBSA		CBSA	
Weight	Ln(HU-Start)		Ln(HU-Start)		Ln(HU-Start)		Ln(HU-Start)		Ln(HU-Start)	
KP F-Stat (91-99, Non-Stack Sample)	36.95		30.94		30.94		29.57		29.87	
MOP F-Stat (91-99, Non-Stack Sample)	36.55		30.87		30.87		29.53		29.76	
KP F-Stat (99-05, Non-Stack Sample)	26.48		25.81		16.95		17.03		17.36	
MOP F-Stat (99-05, Non-Stack Sample)	26.25		25.58		16.78		16.91		17.21	
CoefEqual_Chi2	12.209		8.623		6.813		6.734		6.827	
CoefEqual_PValue	0.000		0.003		0.009		0.009		0.009	

Table 13: **Four Stacked Regressions of House Price Growth on Net Export Growth in Prior (91-99) and Boom (99-05) Periods**

This table reports OLS, reduced-form, first stage and second stages of stacked 2SLS regression $\triangle_{91,99}\&\triangle_{99,05}Ln(HPI_c) = \beta_{91,99} * \triangle_{91,99}\text{NetExp}_m \times Dum_{91,99} + \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \gamma_{91,99} * \boldsymbol{Controls_c} \times Dum_{91,99} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \alpha_{91,99} + \alpha_{99,05} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{91,99}\&\triangle_{99,05}Ln(HPI_c)$ is the stacked growth rate of house price index at county c 91-99 and 99-05. The key variable of interest $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$ are the growth rate of net export at the metropolitan area (CBSA03 code) m 91-99 and 99-05, respectively. $Controls_c$ indicates control variables at county c in the period start year either 1991 or 1999. We use the gravity model-based instrumental variable $\triangle_{91,99}\text{givNetExp}_m$ and $\triangle_{99,05}\text{givNetExp}_m$ as IVs for $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$. Each regression is weighted by the natural logarithm of housing units in the start year (either 1991 or 1999). For the first-stage F-test of two non-stacked samples, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	House Price Growth (91-99 & 99-05, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
Net Export Growth (91-99, An) x Dum91t99	-0.775	-0.207	0.910	1.143	0.989
	(0.995)	(0.936)	(0.910)	(0.925)	(0.916)
Net Export Growth (99-05, An) x Dum99t05	4.508***	3.845***	1.703**	1.984***	2.066***
	(0.994)	(0.926)	(0.691)	(0.692)	(0.724)
R2-adj	0.563	0.645	0.725	0.737	0.747
Panel B. Reduced-form estimates					
GIV Net Export Growth (91-99, An) x Dum91t99	-3.711**	-3.296**	-2.026	-1.707	-1.875
	(1.551)	(1.576)	(1.581)	(1.601)	(1.619)
GIV Net Export Growth (99-05, An) x Dum99t05	7.677***	6.620***	4.660***	5.249***	5.085***
	(1.967)	(1.979)	(1.261)	(1.344)	(1.449)
R2-adj	0.559	0.643	0.728	0.741	0.750
Panel C . 2SLS estimates					
Net Export Growth (91-99, An) x Dum91t99	-3.530**	-3.125*	-2.037	-1.736	-1.902
	(1.684)	(1.687)	(1.719)	(1.743)	(1.767)
Net Export Growth (99-05, An) x Dum99t05	6.926***	6.020***	4.470***	4.976***	4.884***
	(2.327)	(2.316)	(1.725)	(1.805)	(1.780)
Dep Var (Panel D):	Net Export Growth (91-99, An)				
Panel D . First-stage estimates only for 91-99 (non-stack sample)					
GIV Net Export Growth (91-99, An)	1.051***	0.994***	0.994***	0.983***	0.986***
	(0.173)	(0.179)	(0.179)	(0.181)	(0.180)
KP F-Stat	36.95	30.94	30.94	29.57	29.87
MOP F-Stat	36.55	30.87	30.87	29.53	29.76
Dep Var (Panel E):	Net Export Growth (99-05, An)				
Panel E . First-stage estimates only for 99-05 (non-stack sample)					
GIV Net Export Growth (99-05, An)	1.108***	1.100***	1.043***	1.055***	1.041***
	(0.215)	(0.216)	(0.253)	(0.256)	(0.250)
KP F-Stat	26.48	25.81	16.95	17.03	17.36
MOP F-Stat	26.25	25.58	16.78	16.91	17.21
Controls (for all Panels)					
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs (Panel A, B, and C)	1344	1344	1198	1198	1198
Obs (Panel D and E)	672	599	599	599	599
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)

Table 14: **2SLS Stacked Regression of PLMNJ Growth in Prior (91-99) and Boom (99-05) Periods on Net Export Growth (99-05)**

This table reports 2SLS stacked regression $\triangle_{91,99}\&\triangle_{99,05}Ln(PLMNJ_c) = \beta_{91,99} * \triangle_{91,99}\text{NetExp}_m \times Dum_{91,99} + \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \gamma_{91,99} * \boldsymbol{Controls_c} \times Dum_{91,99} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \alpha_{91,99} + \alpha_{99,05} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{91,99}\&\triangle_{99,05}Ln(PLMNJ_c)$ is the stacked growth rate of the dollar amount of private-label mortgages (non-jumbo) (PLMNJ) at county c 91-99 and 99-05. The key variable of interest $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$ are the growth rate of net export at the metropolitan area (CBSA03 code) m 91-99 and 99-05, respectively. $Controls_c$ indicates control variables at county c in the period start year either 1991 or 1999. We use the gravity model-based instrumental variable $\triangle_{91,99}\text{givNetExp}_m$ and $\triangle_{99,05}\text{givNetExp}_m$ as IVs for $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$. Regression is weighted by the natural logarithm of housing units in the start year (either 1991 or 1999). We report the statistics and p-values for the tests of coefficient equality between $\beta_{91,99}$ and $\beta_{99,05}$. For the first-stage F-test of two separate non-stack samples, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS estimates	PLMNJ Growth (91-99 and 99-05, annualized)									
	(1)		(2)		(3)		(4)		(5)	
Net Export Growth (91-99, An) x Dum91t99	-2.112	(4.200)	-2.461	(4.142)	-7.545	(5.144)	-7.796*	(4.562)	-7.784	(4.936)
Net Export Growth (99-05, An) x Dum99t05	14.064***	(3.571)	13.569***	(3.632)	12.188***	(3.503)	12.186***	(3.539)	11.213***	(3.309)
Dum91t99	0.166***	(0.009)	0.339	(0.265)	0.327	(0.368)	0.357	(0.536)	0.409	(0.515)
Dum99t05	0.199***	(0.009)	0.244	(0.219)	0.417*	(0.250)	-0.371	(0.401)	-0.305	(0.408)
Ln(Num of HH, 91) x Dum91t99			0.000	(0.004)	-0.044	(0.051)	-0.107*	(0.058)	-0.113**	(0.049)
Ln(Num of HH, 99) x Dum99t05			0.008***	(0.003)	0.015	(0.034)	-0.007	(0.033)	-0.011	(0.033)
Ln(HH Income, 91) x Dum91t99			-0.013	(0.028)	-0.015	(0.040)	-0.003	(0.056)	-0.002	(0.053)
Ln(HH Income, 99) x Dum99t05			-0.015	(0.024)	-0.035	(0.026)	0.037	(0.039)	0.031	(0.039)
Ratio of Labor Force (1989) x Dum91t99			-0.060	(0.101)	0.002	(0.116)	0.052	(0.101)	-0.040	(0.104)
Ratio of Labor Force (1999) x Dum99t05			0.027	(0.112)	0.103	(0.101)	0.160	(0.107)	0.156	(0.101)
Ln(Num of HU, 91) x Dum91t99					0.044	(0.051)	0.101*	(0.058)	0.108**	(0.049)
Ln(Num of HU, 99) x Dum99t05					-0.003	(0.035)	0.020	(0.034)	0.024	(0.033)
Housing supply elasticity x Dum91t99					-0.010***	(0.003)	-0.007**	(0.003)	-0.004	(0.004)
Housing supply elasticity x Dum99t05					-0.011***	(0.004)	-0.010***	(0.004)	-0.010***	(0.004)
House Vacancy Rate (1989) x Dum91t99					0.267**	(0.113)	0.182*	(0.104)	0.101	(0.098)
House Vacancy Rate (1999) x Dum99t05					0.156	(0.106)	0.164	(0.103)	0.102	(0.101)
Ratio of Renters (1989) x Dum91t99					-0.013	(0.084)	-0.202**	(0.079)	-0.148**	(0.074)
Ratio of Renters (1999) x Dum99t05					-0.109**	(0.055)	-0.056	(0.061)	-0.122**	(0.062)
Ratio of Bachelor Educated (1989) x Dum91t99							-0.078	(0.091)	0.049	(0.105)
Ratio of Bachelor Educated (1999) x Dum99t05							-0.263***	(0.096)	-0.299***	(0.099)
Ratio of White Race (1989) x Dum91t99							-0.084*	(0.046)	-0.066	(0.042)
Ratio of White Race (1999) x Dum99t05							-0.000	(0.033)	-0.023	(0.032)
Ratio of Immigration (80-90) x Dum91t99							1.242***	(0.211)	1.026***	(0.254)
Ratio of Immigration (90-00) x Dum99t05							0.066	(0.191)	0.146	(0.200)
Ratio of Art, Enter, and Recre Emp (1989) x Dum91t99									0.290	(0.473)
Ratio of Art, Enter, and Recre Emp (1999) x Dum99t05									1.774***	(0.643)
Ratio of Health Emp (1989) x Dum91t99									-1.754***	(0.525)
Ratio of Health Emp (1999) x Dum99t05									0.527	(0.444)
Ratio of College Students (1989) x Dum91t99									-0.298	(0.185)
Ratio of College Students (1999) x Dum99t05									0.188	(0.166)
Obs	1404		1404		1248		1248		1248	
Cluster SE	CBSA		CBSA		CBSA		CBSA		CBSA	
Weight	Ln(HU-Start)		Ln(HU-Start)		Ln(HU-Start)		Ln(HU-Start)		Ln(HU-Start)	
KP F-Stat (91-99, Non-stack Sample)	45.44		46.65		34.98		33.97		34.08	
MOP F-Stat (91-99, Non-stack Sample)	45.60		46.92		35.74		34.65		34.74	
KP F-Stat (99-05, Non-stack Sample)	24.68		24.48		16.58		16.85		16.89	
MOP F-Stat (99-05, Non-stack Sample)	23.79		23.62		16.00		16.28		16.36	
CoefEqual_Chi2	9.642		9.165		10.24		12.32		10.54	
CoefEqual_PValue	0.002		0.002		0.001		0.000		0.001	

Table 15: **Four Stacked Regressions of PLMNJ Growth in Prior (91-99) and Boom (99-05) Periods on Net Export Growth (99-05)**

This table reports OLS, reduced-form, first stage and second stages of 2SLS stacked regression $\triangle_{91,99}$ & $\triangle_{99,05}Ln(PLMNJ_c) = \beta_{91,99} * \triangle_{91,99}\text{NetExp}_m \times Dum_{91,99} + \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \gamma_{91,99} * \boldsymbol{Controls_c} \times Dum_{91,99} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \alpha_{91,99} + \alpha_{99,05} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{91,99}\&\triangle_{99,05}Ln(PLMNJ_c)$ is the stacked growth rate of the dollar amount of private-label mortgages (non-jumbo) (PLMNJ) at county c 91-99 and 99-05. The key variable of interest $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$ are the growth rate of net export at the metropolitan area (CBSA03 code) m 91-99 and 99-05, respectively. $Controls_c$ indicates control variables at county c in the period start year either 1991 or 1999. We use the gravity model-based instrumental variable $\triangle_{91,99}\text{givNetExp}_m$ and $\triangle_{99,05}\text{givNetExp}_m$ as IVs for $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$. Each regression is weighted by the natural logarithm of housing units in the start year (either 1991 or 1999). For the first-stage F-test of two non-stack samples, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. We report Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	PLMNJ Growth (91-99 & 99-05, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
Net Export Growth (91-99, An) x Dum91t99	3.771	4.148	1.050	0.214	0.139
	(3.145)	(3.123)	(3.357)	(3.149)	(3.276)
Net Export Growth (99-05, An) x Dum99t05	8.166***	7.858***	6.148***	6.757***	6.466***
	(1.707)	(1.689)	(1.866)	(1.879)	(1.853)
R2-adj	0.819	0.820	0.835	0.844	0.848
Panel B. Reduced-form estimates					
GIV Net Export Growth (91-99, An) x Dum91t99	-2.459	-2.873	-8.112	-8.354*	-8.363*
	(4.798)	(4.728)	(5.040)	(4.541)	(4.918)
GIV Net Export Growth (99-05, An) x Dum99t05	16.612***	15.909***	13.756***	13.769***	12.579***
	(3.473)	(3.565)	(3.561)	(3.741)	(3.746)
R2-adj	0.819	0.820	0.836	0.845	0.849
Panel C . 2SLS estimates					
Net Export Growth (91-99, An) x Dum91t99	-2.112	-2.461	-7.545	-7.796*	-7.784
	(4.200)	(4.142)	(5.144)	(4.562)	(4.936)
Net Export Growth (99-05, An) x Dum99t05	14.064***	13.569***	12.188***	12.186***	11.213***
	(3.571)	(3.632)	(3.503)	(3.539)	(3.309)
Dep Var (Panel D):	Net Export Growth (91-99, An)				
Panel D . First-stage estimates only for 91-99 (non-stack sample)					
GIV Net Export Growth (91-99, An)	1.164***	1.167***	1.075***	1.072***	1.074***
	(0.173)	(0.171)	(0.182)	(0.184)	(0.184)
KP F-Stat	45.44	46.65	34.98	33.97	34.08
MOP F-Stat	45.60	46.92	35.74	34.65	34.74
Dep Var (Panel E):	Net Export Growth (99-05, An) x Dum92t02				
Panel E . First-stage estimates only for 99-05 (non-stack sample)					
GIV Net Export Growth (99-05, An)	1.181***	1.172***	1.129***	1.130***	1.122***
	(0.238)	(0.237)	(0.277)	(0.275)	(0.273)
KP F-Stat	24.68	24.48	16.58	16.85	16.89
MOP F-Stat	23.79	23.62	16.00	16.28	16.36
Controls (for all Panels)					
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs (Panel A, B, and C)	1404	1404	1248	1248	1248
Obs (Panel D and E)	702	702	624	624	624
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)

Table 16: **2SLS Regression of PLMNJ (Non-Owner-Occupied) Growth on PLMNJ (Owner-Occupied) Growth in Boom Period (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{99,05}Ln(PLMNJ_NonOwn_c) = \beta * \triangle_{99,05}Ln(PLMNJ_Own_c) + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05}Ln(PLMNJ_NonOwn_c)$ is the growth rate of the dollar amount of non-owner-occupied private-label mortgages (non-jumbo) at county c 99-05 and the key variable of interest $\triangle_{99,05}Ln(PLMNJ_Own_c)$ is the growth rate of the dollar amount of owner-occupied private-label mortgages (non-jumbo) at county c 99-05. $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05}Ln(PLMNJ_Own_c)$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in 1999 as the weight for the housing market. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	PLMNJ (Non-Owner-Occupied) Growth (07USD, 99-05)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
PLMNJ (Owner-Occupied) Growth (07USD, 99-05,)	1.105***	1.123***	1.124***	1.130***	1.119***
	(0.074)	(0.063)	(0.072)	(0.069)	(0.073)
R2-adj	0.243	0.282	0.262	0.274	0.281
Panel B. Reduced-form estimates					
GIV Net Export Growth (99-05)	16.649***	16.750***	15.038***	15.562***	13.013**
	(6.254)	(5.895)	(5.481)	(5.643)	(5.725)
R2-adj	0.00910	0.0408	0.0439	0.0568	0.0679
Panel C . 2SLS estimates					
PLMNJ (Owner-Occupied) Growth (07USD, 99-05)	1.146***	1.182***	1.307***	1.377***	1.190***
	(0.355)	(0.327)	(0.386)	(0.401)	(0.403)
Dep Var (Panel D):	PLMNJ (Owner-Occupied) Growth (07USD, 99-05)				
Panel D . First-stage estimates					
GIV Net Export Growth (99-05)	14.522***	14.167***	11.509***	11.298***	10.938***
	(3.053)	(3.122)	(3.096)	(3.438)	(3.402)
KP F-Stat	22.63	20.59	13.82	10.80	10.34
MOP F-Stat	21.59	19.77	13.89	10.84	10.28
Controls (for all Panels)					
Basic Controls		Y	Y	Y	Y
Housing Controls			Y	Y	Y
Demographic Controls				Y	Y
Industry Controls					Y
Obs	786	786	695	695	695
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)

Table 17: **2SLS Regression of House Price Growth on Owner-Occupied PLMNJ Growth and Non-Owner-Occupied PLMNJ Growth in Boom Period (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{99,05}Ln(HPI_c) = \beta * \triangle_{99,05}Ln(PLMNJ_Own_c) + \theta *$ Credit-Independent Speculation (c,99-05) $+ \gamma \boldsymbol{Controls_c} + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05}Ln(HPI_c)$ is the growth rate of the house price index at county c 99-05. As a proxy for pure credit expansion, $\triangle_{99,05}Ln(PLMNJ_Own_c)$ is the growth rate of the dollar amount (07USD) of owner-occupied private-label mortgage (non-jumbo) (PLMNJ) at county c 99-05. As a proxy for credit-independent speculation at county c 99-05, Credit-Independent Speculation (c,99-05) is derived from the regression Eq (12), which is a part of growth rate of the non-owner-occupied private-label mortgage (non-jumbo) that cannot be explained by the growth rate of owner-occupied private-label mortgage (non-jumbo). $Controls_c$ indicates control variables at county c in 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05}Ln(PLMNJ_Own_c)$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	House Price Growth (07USD, 99-05)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
PLMNJ Own Growth (Credit Expansion, 99-05)	0.168***	0.156***	0.121***	0.120***	0.115***
	(0.026)	(0.024)	(0.018)	(0.017)	(0.018)
PLMNJ NonOwn Growth (Credit-Ind Speculation, 99-05)	0.011	0.030***	0.017***	0.014***	0.012**
	(0.013)	(0.009)	(0.005)	(0.005)	(0.005)
R2-adj	0.135	0.334	0.504	0.523	0.534
Panel B. Reduced-form estimates					
GIV Net Export Growth (99-05)	7.644***	6.533***	4.635***	5.094***	4.971***
	(1.778)	(1.836)	(1.079)	(1.138)	(1.267)
PLMNJ NonOwn Growth (Credit-Ind Speculation, 99-05)	0.007	0.027***	0.015***	0.012**	0.010**
	(0.013)	(0.009)	(0.006)	(0.005)	(0.005)
R2-adj	0.0514	0.254	0.457	0.481	0.497
Panel C . 2SLS estimates					
PLMNJ Own Growth (Credit Expansion, 99-05)	0.527***	0.461***	0.402***	0.450***	0.454**
	(0.115)	(0.123)	(0.139)	(0.164)	(0.178)
PLMNJ NonOwn Growth (Credit-Ind Speculation, 99-05)	0.019	0.035***	0.022***	0.019**	0.018**
	(0.012)	(0.010)	(0.007)	(0.007)	(0.008)
Dep Var (Panel D):	PLMNJ Own Growth (07USD, 99-05)				
Panel D . First-stage estimates					
GIV Net Export Growth (99-05)	14.508***	14.165***	11.531***	11.332***	10.938***
	(3.055)	(3.123)	(3.091)	(3.431)	(3.398)
PLMNJ NonOwn Growth (Credit-Ind Speculation, 99-05)	-0.022	-0.018	-0.017	-0.016	-0.017
	(0.018)	(0.017)	(0.017)	(0.017)	(0.017)
KP F-Stat	22.55	20.57	13.91	10.91	10.36
MOP F-Stat	21.53	19.76	14.01	10.96	10.30
Controls (for all Panels)					
Basic Controls		Y	Y	Y	Y
Housing Controls			Y	Y	Y
Demographic Controls				Y	Y
Industry Controls					Y
Obs	786	786	786	786	695
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)

Table 18: **2SLS Stacked Regression of Non-Owner-Occupied PLMNJ on Net Export Growth in Prior (91-99) and Boom (99-05) Periods**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{91,99}\&\triangle_{99,05}Ln(PLMNJ_NonOwn_c) = \beta_{91,99} * \triangle_{91,99}\text{NetExp}_m \times Dum_{91,99} + \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \gamma_{91,99} * \boldsymbol{Controls_c} \times Dum_{91,99} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{91,99}\&\triangle_{99,05}Ln(PLMNJ_NonOwn_c)$ is the stacked growth rate of the dollar amount (07USD) of non-owner-occupied private-label mortgages (non-jumbo) (PLMNJ) at county c 91-99 and 99-05. The key variable of interest $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$ are the growth rate of net export at the metropolitan area (CBSA03 code) m 91-99 and 99-05, respectively. $Controls_c$ indicates control variables at county c in the period start year, either 1991 or 1999. We use the gravity model-based instrumental variable $\triangle_{91,99}\text{givNetExp}_m$ and $\triangle_{99,05}\text{givNetExp}_m$ as IVs for $\triangle_{91,99}\text{NetExp}_m$ and $\triangle_{99,05}\text{NetExp}_m$. We report the statistics and p-values for the tests of coefficient equality between $\beta_{91,99}$ and $\beta_{99,05}$. For the first-stage F-test of two separate non-stack samples, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Regression is weighted by the natural logarithm of housing units in the start year (either 1991 or 1999). Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS estimates	Private-label Mortgage (Non-Owner) Growth (91-99 or 99-05, An)				
	(1)	(2)	(3)	(4)	(5)
Panel A. OLS estimates					
Net Export Growth (91-99, An) x Dum91t99	-8.436	-3.039	-6.007	-7.890	-8.312
	(6.412)	(6.272)	(6.873)	(6.749)	(7.250)
Net Export Growth (99-05, An) x Dum99t05	9.939***	10.242***	8.981**	9.478***	9.331**
	(3.764)	(3.509)	(3.722)	(3.642)	(3.647)
R2-adj	0.757	0.772	0.777	0.780	0.784
Panel B. Reduced-form estimates					
GIV Net Export Growth (91-99, An) x Dum91t99	-3.337	-2.596	-3.723	-5.770	-7.115
	(9.855)	(10.204)	(10.714)	(10.836)	(11.911)
GIV Net Export Growth (99-05, An) x Dum99t05	18.950**	18.629***	18.129***	18.251***	17.458**
	(7.458)	(6.941)	(6.946)	(7.008)	(7.026)
R2-adj	0.757	0.771	0.777	0.779	0.783
Panel C . 2SLS estimates					
Net Export Growth (91-99, An) x Dum91t99	-2.894	-2.252	-3.419	-5.315	-6.548
	(8.522)	(8.812)	(9.707)	(9.797)	(10.754)
Net Export Growth (99-05, An) x Dum99t05	16.120**	15.982**	16.129**	16.100***	15.614***
	(6.980)	(6.227)	(6.430)	(6.085)	(5.985)
Dep Var (Panel D):	Net Export Growth (91-99, An)				
Panel D . First-stage estimates only for 91-99 (non-stack sample)					
GIV Net Export Growth (91-99, An)	1.153***	1.153***	1.089***	1.086***	1.087***
	(0.126)	(0.126)	(0.125)	(0.126)	(0.127)
KP F-Stat	83.14	83.86	75.38	74.02	73.77
MOP F-Stat	84.81	85.57	79.14	77.63	77.33
Dep Var (Panel E):	Net Export Growth (99-05, An)				
Panel E . First-stage estimates only for 99-05 (non-stack sample)					
GIV Net Export Growth (99-05, An)	1.181***	1.172***	1.129***	1.130***	1.122***
	(0.238)	(0.237)	(0.277)	(0.275)	(0.273)
KP F-Stat	25.67	25.27	17.22	17.51	17.54
MOP F-Stat	24.94	24.56	16.72	17.06	17.07
Controls (DumPeriod for stacked sample and no DumPeriod for non-stacked sample)					
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs (Panel A, B, and C)	1410	1410	1254	1254	1254
Obs (Panel D and E)	705	705	627	627	627
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)

Table 19: **Four stacked Regressions of Low-minus-High PLMNJ Growth in Boom (99-05) and Bust (05-08) on Net Export Growth (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS stacked regression $\text{LMH}\triangle_{99,05}\&\triangle_{05,08}Ln(PLMNJ_m) = \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \beta_{05,08} * \triangle_{99,05}\text{NetExp}_m \times Dum_{05,08} + \gamma_{99,05} * Dum_{99,05} + \gamma_{05,08} * Dum_{05,08} + \epsilon_{period,m}$. The left-hand-side dependent variable $\text{LMH}\triangle_{99,05}\&\triangle_{05,08}Ln(PLMNJ_m)$ is the differential growth rate of the dollar amount (07USD) of private-label mortgages (non-jumbo) (PLMNJ) at metro m between the low- and high-income ZIP codes 99-05 and 05-08. The dependent variables in column (1) to (4) are the above "low-minus-high" PLMNJ growth based on two, tertile, quartile, and quintile income groups of ZIP codes within metropolitan areas. The key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in metro m in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Low-minus-High PLMNJ Growth (99-05 & 05-08, An, 07USD)			
	(1)	(2)	(3)	(4)
	Half	Tertile	Quartile	Quintile
Panel A. OLS estimates				
Net Export Growth (99-05, An) x Dum99t05	2.679	4.133	4.505*	5.851*
	(2.007)	(2.535)	(2.604)	(3.326)
Net Export Growth (99-05, An) x Dum05t08	-6.226**	-7.816**	-7.194*	-8.051*
	(2.517)	(3.201)	(3.981)	(4.598)
R2-adj	0.195	0.232	0.258	0.247
Panel B. Reduced-form estimates				
GIV Net Export Growth (99-05, An) x Dum99t05	4.037	7.139	10.446*	12.327*
	(3.645)	(4.752)	(5.440)	(6.519)
GIV Net Export Growth (99-05, An) x Dum05t08	-8.614	-18.231***	-16.134***	-21.252***
	(6.379)	(5.739)	(5.895)	(7.560)
R2-adj	0.192	0.237	0.262	0.253
Panel C . 2SLS estimates				
Net Export Growth (99-05, An) x Dum99t05	4.609	8.150	12.074*	14.248**
	(4.137)	(5.351)	(6.292)	(7.020)
Net Export Growth (99-05, An) x Dum05t08	-9.833	-20.813***	-18.648**	-24.563**
	(6.376)	(7.352)	(7.396)	(9.872)
CoefEqual_Chi2	2.803	8.022	7.669	7.567
CoefEqual_PValue	0.094	0.005	0.006	0.006
Dep Var (Panel D):		Net Export Growth (99-05, An)		
Panel D . First-stage estimates only for 99-05 (Non-stack sample)				
GIV Net Export Growth (99-05, An) x Dum99t05	0.876***	0.876***	0.865***	0.865***
	(0.215)	(0.215)	(0.226)	(0.226)
KP F-Stat	16.67	16.67	14.68	14.68
MOP F-Stat	16.39	16.39	14.40	14.40
Obs (Panel A, B, and C)	622	622	596	596
Obs (Panel D)	311	311	298	298
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)

Table 20: **Four Stacked Regressions of Low-minus-High Housing Price Growth in Boom (00-06) and Bust (07-09) Periods on Low-minus-High PLMNJ Growth (99-05)**

This table reports OLS, reduced-form, first stage, and second stages of stacked 2SLS regression $\text{LMH}\triangle_{00,06}$ & $\text{LMH}\triangle_{07,09}Ln(HPI_m) = \beta_{00,06} * \text{LMH}\triangle_{99,05}Ln(PLMNJ_m) \times Dum_{00,06} + \beta_{07,09} * \text{LMH}\triangle_{99,05}Ln(PLMNJ_m) \times Dum_{07,09} + \alpha_{00,06} + \alpha_{07,09} + \epsilon_{period,m}$. The left-hand-side dependent variable $\text{LMH}\triangle_{00,06}$&$\text{LMH}\triangle_{07,09}Ln(HPI_m)$ is the stacked differential growth rate of the house price index between low-income and high-income ZIP codes within metro m 00-06 and 07-09. The key variable of interest $\text{LMH}\triangle_{99,05}Ln(PLMNJ_m$ is the growth rate of the dollar amount of private-label mortgages (non-jumbo) between low-income and high-income ZIP codes within metro m 99-05. We use the gravity model-based instrumental variable $\triangle_{99,05}\text{givNetExp}_m$ as the IV for $\text{LMH}\triangle_{99,05}Ln(PLMNJ_m$. Each regression is weighted by the natural logarithm of housing units in 1999. For the first-stage F-test of the non-stacked sample, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Low-minus-High House Price Growth (00-06 & 07-09, An, 07USD)			
	(1)	(2)	(3)	(4)
	Half	Tertile	Quartile	Quintile
Panel A. OLS estimates				
Low-minus-High PLMNJ Growth (07USD, 99-05, An) x Dum00t06	0.035***	0.032**	0.047***	0.037***
	(0.010)	(0.013)	(0.012)	(0.010)
Low-minus-High PLMNJ Growth (07USD, 99-05, An) x Dum07t09	-0.188***	-0.200***	-0.202***	-0.176***
	(0.037)	(0.044)	(0.048)	(0.039)
R2-adj	0.278	0.292	0.280	0.270
Panel B. Reduced-form estimates				
GIV Net Export Growth (99-05, An) x Dum00t06	0.715**	1.251***	1.544***	1.715***
	(0.330)	(0.430)	(0.538)	(0.582)
GIV Net Export Growth (99-05, An) x Dum07t09	-3.994***	-5.832***	-7.700***	-8.217***
	(1.301)	(1.723)	(2.246)	(2.477)
R2-adj	0.159	0.165	0.175	0.181
Panel C . 2SLS estimates				
Low-minus-High PLMNJ Growth (07USD, 99-05, An) x Dum00t06	0.213	0.222	0.135**	0.127*
	(0.218)	(0.175)	(0.067)	(0.069)
Low-minus-High PLMNJ Growth (07USD, 99-05, An) x Dum07t09	-1.192	-1.036	-0.676**	-0.607**
	(1.066)	(0.675)	(0.268)	(0.271)
CoefEqual_Chi2	1.228	2.248	6.313	4.931
CoefEqual_PValue	0.268	0.134	0.0120	0.0264
Dep Var (Panel D):	Low-minus-High PLMNJ Growth (07 USD, 99-05, An)			
Panel D . First-stage estimates only for 00-06 (Non-stack sample)				
GIV Net Export Growth (99-05, An)	3.351	5.630	11.397**	13.532**
	(3.186)	(4.079)	(4.705)	(6.251)
KP F-Stat	1.106	1.905	5.867	4.685
MOP F-Stat	1.068	1.823	5.559	4.374
Obs (Panel A, B, and C)	582	582	532	532
Obs (Panel D)	291	291	266	266
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)

Table 21: **Four stacked Regressions of Low-minus-High PLMNJ Number Growth in Boom (99-05) and Bust (05-08) on Net Export Growth (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS stacked regression LMH$\triangle_{99,05}$&$\triangle_{05,08}Ln(PLMNJ_m) = \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \beta_{05,08} * \triangle_{99,05}\text{NetExp}_m \times Dum_{05,08} + +\gamma_{99,05} * Dum_{99,05} + \gamma_{05,08} * Dum_{05,08} + \epsilon_{period,m}$. The left-hand-side dependent variable LMH$\triangle_{99,05}$&$\triangle_{05,08}Ln(PLMNJ_m)$ is the differential growth rate of the number of private-label mortgages (non-jumbo) (PLMNJ) at metro m between the low- and high-income ZIP codes 99-05 and 05-08. The dependent variables in columns (1) to (4) are the above "low-minus-high" PLMNJ growth based on two, tertile, quartile, and quintile income groups of ZIP codes within metropolitan areas. The key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in metro m in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Low-minus-High PLMNJ Number Growth (99-05 & 05-08, An, 07USD)			
	(1)	(2)	(3)	(4)
	Half	Tertile	Quartile	Quintile
Panel A. OLS estimates				
Net Export Growth (99-05, An) x Dum99t05	1.867	3.634	3.295	5.110
	(2.049)	(2.513)	(2.509)	(3.167)
Net Export Growth (99-05, An) x Dum05t08	-5.126**	-6.758**	-6.181*	-7.423*
	(2.548)	(2.896)	(3.537)	(4.254)
R2-adj	0.169	0.196	0.214	0.207
Panel B. Reduced-form estimates				
GIV Net Export Growth (99-05, An) x Dum99t05	3.923	8.176*	10.919**	14.132**
	(3.643)	(4.458)	(5.072)	(5.926)
GIV Net Export Growth (99-05, An) x Dum05t08	-5.410	-15.923***	-15.482***	-20.854***
	(6.455)	(5.248)	(5.523)	(7.185)
R2-adj	0.166	0.202	0.221	0.216
Panel C . 2SLS estimates				
Net Export Growth (99-05, An) x Dum99t05	4.479	8.150	12.620**	16.333**
	(4.217)	(5.431)	(6.406)	(6.913)
Net Export Growth (99-05, An) x Dum05t08	-6.176	-18.178***	-17.894**	-24.103**
	(6.705)	(6.692)	(7.093)	(9.529)
CoefEqual_Chi2	1.378	6.973	6.757	7.568
CoefEqual_PValue	0.240	0.008	0.009	0.006
Dep Var (Panel D):	Net Export Growth (99-05, An)			
Panel D . First-stage estimates only for 99-05 (Non-stack sample)				
GIV Net Export Growth (99-05, An) x Dum99t05	0.876***	0.876***	0.865***	0.865***
	(0.215)	(0.215)	(0.226)	(0.226)
KP F-Stat	16.67	16.67	14.68	14.68
MOP F-Stat	16.39	16.39	14.40	14.40
Obs (Panel A, B, and C)	622	622	596	596
Obs (Panel D)	311	311	298	298
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)

Table 22: **Four stacked Regressions of Low-minus-High PLMNJ (Owner-Occupied) Growth in Boom (99-05) and Bust (05-08) on Net Export Growth (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS stacked regression $\text{LMH}\triangle_{99,05}\&\triangle_{05,08}Ln(PLMNJ_Own_m) = \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \beta_{05,08} * \triangle_{99,05}\text{NetExp}_m \times Dum_{05,08} + +\gamma_{99,05} * Dum_{99,05} + \gamma_{05,08} * Dum_{05,08} + \epsilon_{period,m}$. The left-hand-side dependent variable $\text{LMH}\triangle_{99,05}\&\triangle_{05,08}Ln(PLMNJ_Own_m)$ is the differential growth rate of the dollar amount (07USD) of private-label mortgages (non-jumbo) (PLMNJ) (Owner-Occupied) at metro m between the low- and high-income ZIP codes 99-05 and 05-08. The dependent variables in column (1) to (4) are the above "low-minus-high" PLMNJ growth based on two, tertile, quartile, and quintile income groups of ZIP codes within metropolitan areas. The key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in metro m in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Low-minus-High PLMNJ (Owner) Growth (99-05 & 05-08, An, 07USD)			
	(1)	(2)	(3)	(4)
	Half	Tertile	Quartile	Quintile
Panel A. OLS estimates				
Net Export Growth (99-05, An) x Dum99t05	2.360	4.038	4.505*	5.739*
	(2.036)	(2.554)	(2.599)	(3.330)
Net Export Growth (99-05, An) x Dum05t08	-5.566**	-6.747**	-6.266	-7.636
	(2.397)	(3.061)	(4.029)	(4.715)
R2-adj	0.150	0.195	0.220	0.201
Panel B. Reduced-form estimates				
GIV Net Export Growth (99-05, An) x Dum99t05	3.477	6.941	10.170*	12.661**
	(3.689)	(4.640)	(5.205)	(6.355)
GIV Net Export Growth (99-05, An) x Dum05t08	-7.266	-16.524***	-15.170***	-19.566**
	(6.194)	(5.259)	(5.745)	(7.568)
R2-adj	0.147	0.200	0.225	0.206
Panel C . 2SLS estimates				
Net Export Growth (99-05, An) x Dum99t05	3.970	7.924	11.754*	14.633**
	(4.164)	(5.306)	(6.148)	(6.959)
Net Export Growth (99-05, An) x Dum05t08	-8.295	-18.864***	-17.533**	-22.614**
	(6.233)	(6.922)	(7.294)	(9.792)
CoefEqual_Chi2	1.996	7.150	7.316	7.185
CoefEqual_PValue	0.158	0.007	0.007	0.007
Dep Var (Panel D):	Net Export Growth (99-05, An)			
Panel D . First-stage estimates only for 99-05 (Non-stack sample)				
GIV Net Export Growth (99-05, An) x Dum99t05	0.876***	0.876***	0.865***	0.865***
	(0.215)	(0.215)	(0.226)	(0.226)
KP F-Stat	16.67	16.67	14.68	14.68
MOP F-Stat	16.39	16.39	14.40	14.40
Obs (Panel A, B, and C)	622	622	596	596
Obs (Panel D)	311	311	298	298
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)

Table 23: **Four stacked Regressions of Low-minus-High GSEM Growth in Boom (99-05) and Bust (05-08) on Net Export Growth (99-05)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS stacked regression $\text{LMH}\triangle_{99,05}\&\triangle_{05,08}Ln(GSEM_m) = \beta_{99,05} * \triangle_{99,05}\text{NetExp}_m \times Dum_{99,05} + \beta_{05,08} * \triangle_{99,05}\text{NetExp}_m \times Dum_{05,08} + \gamma_{99,05} * Dum_{99,05} + \gamma_{05,08} * Dum_{05,08} + \epsilon_{period,m}$. The left-hand-side dependent variable $\text{LMH}\triangle_{99,05}\&\triangle_{05,08}Ln(GSEM_m)$ is the differential growth rate of the dollar amount (07USD) of government-sponsored enterprise mortgages (GSEM) at metro m between the low- and high-income ZIP codes 99-05 and 05-08. The dependent variables in column (1) to (4) are the above "low-minus-high" GSEM growth based on two, tertile, quartile, and quintile income groups of ZIP codes within metropolitan areas. The key variable of interest $\triangle_{99,05}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 99-05. We use the gravity model-based instrumental variable ($\triangle_{99,05}\text{givNetExp}_m$) as IV for $\triangle_{99,05}\text{NetExp}_m$. For the first-stage F-test, we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in metro m in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Low-minus-High GSEM Growth (99-05 & 05-08, An, 07USD)			
	(1)	(2)	(3)	(4)
	Half	Tertile	Quartile	Quintile
Panel A. OLS estimates				
Net Export Growth (99-05, An) x Dum99t05	0.268	1.268	2.774	1.543
	(1.738)	(2.301)	(2.501)	(2.753)
Net Export Growth (99-05, An) x Dum05t08	-4.217**	-3.143	-4.371	-3.972
	(2.125)	(2.677)	(3.088)	(3.455)
R2-adj	0.00631	0.0137	0.0318	0.0227
Panel B. Reduced-form estimates				
GIV Net Export Growth (99-05, An) x Dum99t05	-4.208	-3.735	-0.013	0.292
	(2.866)	(3.731)	(3.901)	(4.841)
GIV Net Export Growth (99-05, An) x Dum05t08	0.313	1.167	-1.137	-3.740
	(4.919)	(4.536)	(5.766)	(6.925)
R2-adj	0.00226	0.0125	0.0272	0.0213
Panel C . 2SLS estimates				
Net Export Growth (99-05, An) x Dum99t05	-4.804	-4.264	-0.015	0.337
	(3.501)	(4.477)	(4.491)	(5.564)
Net Export Growth (99-05, An) x Dum05t08	0.357	1.332	-1.314	-4.322
	(5.642)	(5.307)	(6.482)	(7.532)
CoefEqual_Chi2	0.457	0.432	0.019	0.175
CoefEqual_PValue	0.499	0.511	0.889	0.675
Dep Var (Panel D):	Net Export Growth (99-05, An)			
Panel D . First-stage estimates only for 99-05 (Non-stack sample)				
GIV Net Export Growth (99-05, An) x Dum99t05	0.876***	0.876***	0.865***	0.865***
	(0.215)	(0.215)	(0.226)	(0.226)
KP F-Stat	16.67	16.67	14.68	14.68
MOP F-Stat	16.39	16.39	14.40	14.40
Obs (Panel A, B, and C)	622	622	596	596
Obs (Panel D)	311	311	298	298
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)	Ln(CBSAHU99)

A Appendix

A.1 Appendix for Data Details

A.1.1 HMDA Mortgage Data Processing Procedures at ZIP level

In this section, I illustrate the processing details regarding the mortgage data at the ZIP code level. I will describe the processes of crosswalking from census tracts to ZIP codes, retaining data in the 1990-2000-unchanged ZIP codes, linking ZIP codes to primary counties and metropolitan areas, mortgage data requirement at the ZIP level, and grouping ZIP codes into income groups within metropolitan areas.

A.1.1.1 Crosswalk from Census Tracts to ZIP Codes

First, I use the 712 or 800 consistent counties across time to filter HMDA data, depending on the period of interest. Second, HMDA mortgage data uses various versions of census tracts across time. 1981-1991, 1992-2002, and 2003-2009 HMDA datasets use 1980, 1990, and 2000 versions of the census tract, respectively. From the Missouri Census Data Center [65], I obtain the crosswalk file that links 1980 census tracts to 1990 ZIP codes, that links 1990 census tracts to 1990 ZIP codes, and that links 2000 census tracts to 2000 ZIP codes. I use population weight in the crosswalk since multiple census tracts can match multiple ZIP codes. I use 1992-1999 as the prior period for comparison because I intend to retain a reasonably large number of ZIP codes in a balanced panel for comparison. To prevent within-ZIP code changes from contaminating my analysis, I only keep ZIP codes that are mostly covered by HMDA reporting. Specifically, I keep ZIP codes with populations crosswalked from the HMDA mortgage data that are equal to or above 90% of the ZIP total population. The deleted ZIP codes are mostly the ZIP codes on the boundary of metropolitan areas, since HMDA mainly covers metropolitan areas.

ZIP codes are reasonably large areas within metropolitan areas. One ZIP code typically contains several census blocks, which are designed to provide more statistical uniformity with an average of 4,000 plus population. In the 1992-2009 mortgage panel data of 9995 ZIP codes, ZIP codes contain a population size of 4237, 12610, and 26023 at 25%, 50%, and 75%, respectively. [66]

A.1.1.2 Retaining Data in the 1990-2000-unchanged ZIP codes

According to Bailey and Helmuth (2023), ZIP code geographic boundary can change in relationship to population growth and decline across time. For example, their detailed work based on The Postal

[65] Missouri Census Data Center's website is here: `https://mcdc.missouri.edu/applications/geocorr.html`

[66] For more information comparing census tracts and ZIP codes, see `https://proximityone.com/tracts_zips.htm`.

Bulletin by the U.S. Postal Office shows that 3015 changes happened between 1990 and 2000. This number of changes is not small given 29,401 distinct ZIP codes were in operation at the beginning of 1990 (by the crosswalk file from the Missouri Census Data Center described above). Typical changes include merge, split, and boundary alteration. These changes can result in measurement error (matching observations to the wrong unites) or missing data (due to an observation reporting a ZIP code that did not in operation at the beginning of the sample period).

To avoid the potential measurement error and missing data, I get a list of 1990-2000-unchanged ZIP codes by following steps. First, from the Missouri Census Data Center, I get the crosswalk file from 1990 ZIP codes to 1990 counties with population weights. Second, I delete any ZIP codes that may experience change during 1990-2000 from the 1990 ZIP codes list. Specifically, I delete ZIP codes that appear in the following variables in the following files by Bailey and Helmuth (2023): (1) variable "oldzipcode" or "newzipcode" in file "ZIP_Code_Changes_1990-2000_Documentation.xls" and (2) "zipcode" in file "ZIP_xwalk_90-00.dta". Out of the 29,401 ZIP codes in the year 1990, 26,577 ZIP codes did not experience any change. Third, I match these ZIP codes to ZIP codes in the crosswalk file from 2000 ZIP codes to 2000 counties with population weights. I retain only matched ZIP codes. The above three steps result in 26,058 1990-2000-unchanged ZIP codes.

A.1.1.3 Link ZIP codes to Their Primary Counties and Metropolitan Areas

One ZIP code can be linked to multiple counties (metropolitan areas) when it exists on the border. For each match zip code, I only keep its primary matched county that contains the most population in the zip code. From the Missouri Census Data Center, I get the crosswalk file that links 2000 ZIP codes to 2000 counties. I retain counties in the 2003 version of metropolitan areas in U.S. mainland (excluding Hawaii and Alaska).[67]

The above steps result in 13,513 1990-2000-unchanged ZIP codes that are matched to primary counties in the metropolitan areas (2003 version CBSA code) in the U.S. mainland. I use these ZIP codes to filter HMDA ZIP-level data.

A.1.1.4 Mortgage Data Requirement at the ZIP Level

I require each ZIP code to have mortgage data in all years of analysis (1992-2009 for figure and 1999-2009 for regression analysis). I also require each ZIP code must have nonzero value for government-sponsored enterprise mortgages, private-label mortgages (non-jumbo), and mortgages retained on the balance sheet of financial institutions in the initial year (either 1992 or 1999). This requirement make sure that each ZIP code has starting value and non-missing growth rate in the

[67] Metropolitan areas definition (CBSA code) of 2003 can be found here: `https://www2.census.gov/programs%2Dsurveys/metro%2Dmicro/geographies/reference%2Dfiles/2003/historical%2Ddelineation%2Dfiles/030606omb%2Dnecta%2Dcnecta.xls`.

main analysis.[68] The resulting 1992-2009 (1999-2009) mortgage panel data contains 9,995 (10,785) ZIP codes from 693 (782) counties.

A.1.1.5 Group ZIP codes Within Metropolitan Areas

To conduct analysis in Section (7) for the model-based new predictions, I need to sum mortgage data in ZIP codes into half-, tertile-, quartile-, and quintile-income groups within metropolitan areas. To make sure each metropolitan statistic area (MSA) contains enough ZIP codes for calculation, I require that each MSA contains at least three ZIP codes for the calculation of half- and tertile-income groups and at least five ZIP codes for the calculation of quartile- and quintile-income groups. Since ZIP codes are reasonably large areas (with a population size of 4237 and 26023 at 25% and 75% in the 1992-2009 mortgage panel data), I believe three (five) ZIP codes are a good minimum number for the calculation. For descriptive figures between 1992 and 2009, we have 9,969 ZIP codes from 296 metros (688 counties) for half- and tertile-income groups and 9,917 ZIP codes from 281 metros (671 counties) for quartile- and quintile-income groups. For regression analysis between 1999 and 2009, we have 10,758 ZIP codes from 314 metros (776 counties) for half- and tertile-income groups and 10,710 ZIP codes from 300 metros (761 counties) for quartile- and quintile-income groups.

A.1.2 House Price Data at ZIP level

The annual house price index at the ZIP level based on repeat sales comes from the Federal Housing Financing Agency.[69] This database has reasonably good coverage of the ZIP codes of US metropolitan areas in and after 1992 (11,556 ZIP codes in 1992) but a little bit less coverage than the HMDA mortgage data.

A.1.2.1 Merge House Price and Mortgage Data at ZIP Level

For both figure and regression analysis regarding house prices, I require that the ZIP codes are covered by both house price data and mortgage data refined in Section A.1.1.4. The merged data set contains fewer ZIP codes compared to mortgage data since house price data covers fewer ZIP codes. Second, I require that each zip code contains both data 1992-2009 (1999-2009) for figure (regression analysis).

[68] This step also excludes ZIP codes with a tiny population.

[69] The data website of the Federal Housing Financing Agency is here: `https://www.fhfa.gov/DataTools/Downloads/Pages/House-Price-Index-Datasets.aspx#qexe`.

A.1.2.2 Interpolating House Units

I collect the number of house units at the ZIP level from Decennial Census Summary Files in 1900 (File 3B), 2000 (File 1), and 2010 (File 1). Since Census data are collected as of March in 1990, 2000, and 2010, I treat the number of house units as the one at the end of last year (1989, 1999, and 2009, respectively). I interpolate the number of house units from 1990 to 2008 by a linear function. Then, I merge house unit data with the prices-mortgage merged data. For the final data, I require that each MSA contains at least three ZIP codes for the calculation of half- and tertile-income groups and at least five ZIP codes for the calculation of quartile- and quintile-income groups.

Due to data availability of house prices and house units, the final data contains fewer ZIP codes than the mortgage data. For descriptive figures between 1992 and 2009, we have 6,083 (5,942) ZIP codes from 244 (204) metropolitan areas for half- (quintile-) income groups. For regression analysis between 1999 and 2009, we have 8,517 (8,427) ZIP codes from 293 (268) metropolitan areas for half- (quintile-) income group.

Six Metropolitan Areas Severely Affected by 2005 Hurricanes Following Section 3.2, we delete six metropolitan areas in our regression analysis that contain the twelve counties that were severely affected by 2005 Hurricanes. These six metropolitan areas are Abbeville (LA, CBSA03 code: 10020), Gulfport-Biloxi (MS, 25060), Key West-Marathon (FL, 28580), Lake Charles (LA, 29340), New Orleans-Metairie-Kenner (LA, 35380), and Pascagoula (MS, 37700).

A.2 Gravity Model-based IV: US Imports

I have illustrated the key idea of the gravity model-based instrument for US exports by Feenstra, Ma, and Xu (2019) in Section 2.2. For completeness, I show how they construct IV for US imports here. The gravity-based IV for US imports starts from a simple symmetric constant-elasticity equation by Romalis (2007):

$$\frac{X_{s,v,t}^{j,US}}{X_{s,v,t}^{j,i}} = \left(\frac{w_{s,t}^{j} d^{j,US} \tau_{s,t}^{j,US}}{w_{s,t}^{j} d^{j,i} \tau_{s,t}^{j,i}}\right)^{1-\sigma} \frac{(P_{s,t}^{US})^{\sigma-1} E_{s,t}^{US}}{(P_{s,t}^{i})^{\sigma-1} E_{s,t}^{i}} = \left(\frac{d^{j,US} \tau_{s,t}^{j,US}}{d^{j,i} \tau_{s,t}^{j,i}}\right)^{1-\sigma} \frac{(P_{s,t}^{US})^{\sigma-1} E_{s,t}^{US}}{(P_{s,t}^{i})^{\sigma-1} E_{s,t}^{i}} \tag{28}$$

In the above formula, $X_{s,v,t}^{j,US}$ is country j's export to US in product variant v in industry s in year t. $X_{s,v,t}^{j,i}$ is the similar term but representing country j's export to country i. $w_{s,t}^{j}$ is the relative marginal cost of production in industry s in country j, which is canceled out in the above equation. $\tau_{s,t}^{j,US}$ and $\tau_{s,t}^{j,i}$ are the *ad valorem* total import tariff on country j's export to the US and country i. $d^{j,US}$ and $d^{j,i}$ are the bilateral distance and other fixed trade costs from country j to the US and to country i. $P_{s,t}^{US}$ and $P_{s,t}^{i}$ are the aggregate price index in the US and country i; $E_{s,t}^{US}$ and $E_{s,t}^{i}$ are the total expenditure in the US and country i. Lastly, σ denotes the constant elasticity of substitution ($\sigma > 1$).

Like the one for export, the intuition of this gravity-style model for import is quite straightfor-

ward. The ratio of country i's export to the US relative to country j is declining with the ratio of bilateral distance and the ratio of *ad valorem* import tariff, but rising with the ratio of aggregate price index and total expenditure.

Assume that there are $N^j_{s,t}$ identical product varieties exported by country j in industry s and year t, one can re-arrange the above equation, multiply both sides with $N^j_{s,t}$, and sum over countries $i \neq US$:

$$N^j_{s,t}X^{j,US}_{s,v,t} * \sum_{i\neq US}\left[(d^{j,i})^{1-\sigma}(P^i_{s,t})^{\sigma-1}E^i_{s,t}\right] = (d^{j,US}\tau^{j,US}_{s,t})^{1-\sigma}(P^{US}_{s,t})^{\sigma-1}E^{US}_{s,t} * \sum_{i\neq US}\left[N^j_{s,t}X^{j,i}_{s,v,t}(\tau^{j,i}_{s,t})^{\sigma-1}\right]$$

Like before, as the above equation holds for any countries $i \neq US$, one can choose the set of countries that have similar economic conditions with the US (so that they are buyers competing with US buyers when country j considers its export) to make my prediction more accurate. Feenstra, Ma, and Xu (2019) use the eight high-income countries proposed by Autor, Dorn, and Hanson (2013b).

I denote the sectoral export by country j to the US and to country i as $X^{j,US}_{s,t} \equiv X^{j,US}_{s,v,t} * N^j_{s,t}$ and $X^{j,i}_{s,t} \equiv X^{j,i}_{s,v,t} * N^j_{s,t}$. Then I can get

$$X^{j,US}_{s,t} * \sum_{i\neq US}\left[(d^{j,i})^{1-\sigma}(P^i_{s,t})^{\sigma-1}E^i_{s,t}\right] = (d^{j,US}\tau^{j,US}_{s,t})^{1-\sigma}(P^{US}_{s,t})^{\sigma-1}E^{US}_{s,t} * \sum_{i\neq US}\left[X^{j,i}_{s,t}(\tau^{j,i}_{s,t})^{\sigma-1}\right]$$

After a few re-arrangement, I can get the formula for $X^{j,US}_{s,t}$:

$$X^{j,US}_{s,t} = \frac{(d^{j,US}\tau^{j,US}_{s,t})^{1-\sigma}(P^{US}_{s,t})^{\sigma-1}E^{US}_{s,t}}{\sum_{i\neq US}\left[(d^{j,i})^{1-\sigma}(P^i_{s,t})^{\sigma-1}E^i_{s,t}\right]} * \left(\sum_{k\neq US}X^{j,k}_{s,t}\right) * \left\{\sum_{i\neq US}\left[\frac{X^{j,i}_{s,t}}{\sum_{k\neq US}X^{j,k}_{s,t}}(\tau^{j,i}_{s,t})^{\sigma-1}\right]\right\} \quad (29)$$

Note in the above formula, I multiply and divide by $\sum_{k\neq US}X^{j,k}_{s,t}$ to prepare for the regression setup in the next step. Now I can take logs of the above equation and move the term $\ln\left(\sum_{k\neq US}X^{j,k}_{s,t}\right)$ to the left-hand side of the equation to get the regression-style formula:

$$\begin{aligned}\ln\left(X^{j,US}_{s,t}\right) = &\underbrace{\ln\left(\sum_{k\neq US}X^{j,k}_{s,t}\right)}_{\text{Term 0}} + \underbrace{\ln\left((P^{US}_{s,t})^{\sigma-1}E^{US}_{s,t}\right)}_{\text{Ind-Year FE: }\gamma^{US}_{s,t}} + \underbrace{(1-\sigma)\ln\left(d^{j,US}\right)}_{\text{Exporting-country FE: }\delta^{j,US}} \\ &+ \underbrace{(1-\sigma)\ln\left(\tau^{j,US}_{s,t}\right)}_{\text{Term 1}} + \underbrace{(\sigma-1)\ln\left(\left\{\sum_{i\neq US}\left[\frac{X^{j,i}_{s,t}}{\sum_{k\neq US}X^{j,k}_{s,t}}(\tau^{j,i}_{s,t})^{\sigma-1}\right]\right\}^{\frac{1}{\sigma-1}}\right)}_{\text{Term 2: }(\sigma-1)\ln\left(T^j_{s,t}\right)} + \epsilon^j_{s,t}\end{aligned} \quad (30)$$

Now, the US import from country j in the industry s year t can be decomposed into six terms. "Term 0" is the other eight high-income countries' import from country j, which reflects the world supply. The second term $\gamma^{US}_{s,t}$ reflects the US demand shocks, which is potentially endogenous. I remove this term by the US industry-by-year fixed effects. The third term $\delta^{j,US}$ reflects the distance from country j to the US and all other industry- and year-invariant trade costs. Since this term is predetermined rather than a shock, I remove it by the exporting-country fixed effects.

"Term 1" is the tariff on country j's exports imposed by the US, which is out of the control of exporting firms. I keep this term to capture the shock from tariffs. "Term 2" is the weighted average tariffs on country j's exports imposed by other eight high-income countries. Intuitively, when this weighted average tariffs on country j's exports rise, destination country j will export to the US as a substitution. I keep this term to capture this substitution effect. The last term $\epsilon^j_{s,t} = -\ln\left(\sum_{i\neq US}[(d^{j,i})^{1-\sigma}(P^i_{s,t})^{\sigma-1}E^i_{s,t}]\right)$ is unobserved and remains in the regression error term.

After the above regression, I can isolate predicted US imports that are presumably exogenous:

$$\ln\left(\widehat{X^{j,US}_{s,t}}\right) = \ln\left(\sum_{k\neq US} X^{j,k}_{s,t}\right) + \hat{\beta}_1 * \ln\left(\tau^{j,US}_{s,t}\right) + \hat{\beta}_2 * \ln\left(T^j_{s,t}\right) \tag{31}$$

A.3 Appendix for the Empirical Results Supporting Credit Expansion

A.3.1 Exclusion Restriction: Test on House Price Growth

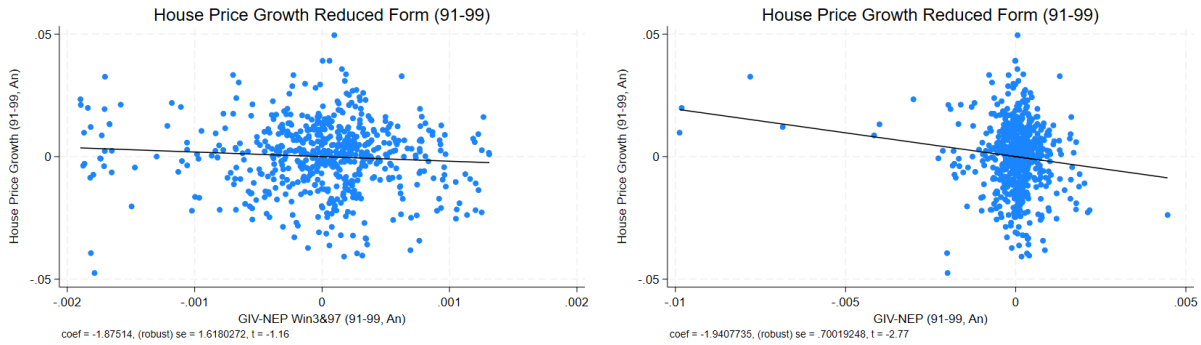
In section 5.2 Table (12) and (13), in which I use prior period to test the exclusion restriction of my instrumental variable approach, I winsorize net export growth and its GIV in the prior period (1991-1999). In this appendix, I will compare results with and without winsorization to show that a few outliers can make the coefficient of net export growth significantly negative.

First, for the prior period (1991-1999), the 2SLS with and without winsorization can be seen from the tiny table below. We can see that 2SLS estimates without winsorization is -1.809 and significant at the 1% level. However, 2SLS estimates with winsorization at the 3% and 97% level is -1.902 but not significant. Comparison between two results show that a few ourliers can make the impact of net export growth on house price growth significantly negative.

Table A.1: **2SLS Regression of House Price Growth on Net Export Growth in Prior (91-99) With and Without Winsorization**

TSLS estimates	House Price Growth (07USD, 91-99, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. Sample with Wonsorization					
Net Export Growth (91-99, An) x Dum91t99	-3.530**	-3.125*	-2.037	-1.736	-1.902
	(1.684)	(1.687)	(1.719)	(1.743)	(1.767)
Panel B. Sample without Wonsorization					
Net Export Growth (91-99, An)	-2.295***	-2.142***	-1.854**	-1.759**	-1.809***
	(0.700)	(0.726)	(0.722)	(0.706)	(0.701)
Controls					
Basic Controls		Y	Y	Y	Y
Housing Controls			Y	Y	Y
Demographic Controls				Y	Y
Industry Controls					Y
Obs for Both Panels	672	672	599	599	599
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)

Another way to illustrate the above effect of a few outliers is to compare the added-variable plots (AV-plot) of two reduced-form regressions. In Figure (A.1), the left subfigure shows the AV-plot of the reduced-form regression for the sample with winsorization and the right subfigure is for the sample without winsorization. Without winsorization, a few outliers make the impact of net export growth on house price grow significantly negative in the right subfigure. By winsorization, the left subfigure shows that, for most observations, such negative is insignificant.



(a) With Winsorization (b) Without Winsorization

Figure A.1: Reduced-form Regression of House Price Growth on Net Export Growth in Prior Period (1991-1999) With and Without Winsorization at 3% and 97%

A.3.2 Exclusion Restriction: Test on PLMNJ Growth

In section 5.2 Table (14) and (15), in which I use the prior period to provide suggestive validation for the exclusion restriction of my instrumental variable approach, I drop three outliers based on the GIV of net export growth in the prior period (1991-1999). These three outliers are Howard County (county FIPS: 18067), Tipton County (18159), and Durham County (37063). In this appendix, I will compare results with and without dropping outliers to show that a few outliers can make the coefficient of net export growth significantly negative.

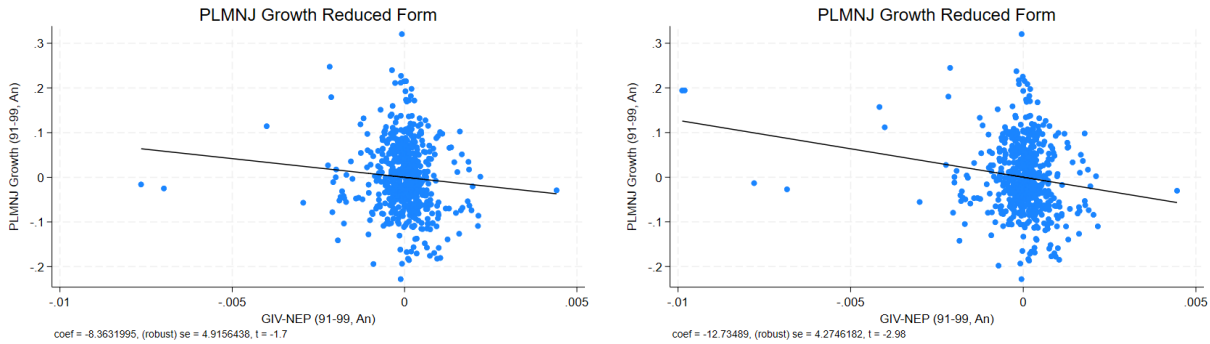
First, for the prior period (1991-1999), the 2SLS with and without dropping three outliers can be seen from the tiny table below. We can see that 2SLS estimates without dropping is -11.720 and significant at the 1% level. However, 2SLS estimates with dropping the three outliers is -7.784 but not significant. Comparison between two results show that a few outliers can make the impact of net export growth on private-label mortgages (non-jumbo) growth significantly negative.

Another way to illustrate the above effect of a few outliers is to compare the added-variable plots (AV-plot) of two reduced-form regressions. In Figure (A.2), the left subfigure shows the AV-plot of the reduced-form regression for the sample with dropping the three outliers and the right subfigure is for the sample without dropping outliers. Without dropping, a few outliers make the impact of net export growth on private-label mortgage grow significantly negative in the right subfigure.

Table A.2: **2SLS Regression of Private-label Mortgage (non-jumbo) Growth on Net Export Growth in Prior (91-99) With and Without Dropping Outliers**

TSLS estimates	Private-label Mortgage (non-jumbo) Growth (07USD, 91-99, annualized)				
	(1)	(2)	(3)	(4)	(5)
Panel A. Sample with Dropping Three Outliers					
Net Export Growth (91-99, An)	-2.112	-2.461	-7.545	-7.796*	-7.784
	(4.200)	(4.142)	(5.144)	(4.562)	(4.936)
Panel B. Sample without Dropping Three Outliers					
Net Export Growth (91-99, An)	-6.108	-6.404	-11.185**	-11.351***	-11.720***
	(4.297)	(4.226)	(4.538)	(4.055)	(4.308)
Controls					
Basic Controls		Y	Y	Y	Y
Housing Controls			Y	Y	Y
Demographic Controls				Y	Y
Industry Controls					Y
Obs for Panel A	702	702	624	624	624
Obs for Panel B	705	705	627	627	627
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)	Ln(HU-Start)

Figure A.2: Reduced-form Regression of Private-label Mortgage (non-jumbo) Growth on Net Export Growth in Prior Period (1991-1999) With and Without Dropping the Three Outliers



(a) With Dropping Outliers

(b) Without Dropping Outliers

By dropping the three outliers, the left subfigure shows that, for most observations, such negative impact is insignificant.

A.4 Robustness: State Heterogeneity

In this subsection, I perform robustness tests to support the main conclusion that credit expansion in private-label mortgages induced by net export growth causes the 1999-2009 U.S. house price boom and bust across metropolitan areas. I show that this main conclusion is robust to state-level differences in anti-predatory lending laws (Di Maggio and Kermani, 2017), recourse laws (Ghent and Kudlyak, 2011), judicial requirement in foreclosure (Mian, Sufi, and Trebbi, 2015), distinction between sand states and other states (Choi, Hong, Kubik, and Thompson, 2016), and state capital

gain tax (Gao, Sockin, and Xiong, 2020).

A.4.1 Anti-Predatory-Lending States vs. Other States

In this subsection, my main conclusion holds even without the cross-state difference in anti-predatory lending laws. In addition, I show that growth in private-label mortgages leads to a stronger housing boom and a stronger bust in states with anti-predatory lending laws after 2004.

According to Di Maggio and Kermani (2017), prior to 2004, a dozen of states had already implemented anti-predatory laws to protect mortgage borrowers against unfair and deceptive practices. However, on January 7th, 2004, the Office of the Comptroller of the Currency (OCC) preempted national banks (rather than state-chartered depository institutions and independent mortgage companies) from state antipredatory-lending law (APL law). They show that such deregulation results in the national bank's credit expansion (relative to state-chartered depository institutions and independent mortgage companies), house price growth, and nontradable employment rise in 2004-2006, but a sharp decline subsequently in these states relative to other states.[70] Thus, in the period 2004-2005, all states that have implemented APL laws before cannot apply such laws to national banks. Appendix Table 1 from Di Maggio and Kermani (2017) summarizes the list of APL states before 2004. Since I only have annual HMDA data, I restrict my APL states to the states that implemented APL at least half a year before 2004, resulting in eleven APL states.[71]

My stacked 2SLS regression is

$$\begin{aligned}\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = {} & \beta_{Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} \\ & + \beta_{APL,Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} \times Dum_{APL} + \beta_{APL,Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} \times Dum_{APL} \\ & + \gamma_{Boom} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{Bust} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_c\end{aligned} \tag{32}$$

Controls, weight, and standard errors are the same as Eq(9).

Table (A.3) reports the above 2SLS results and shows two key conclusion. First, compared to non-APL states, APL-states experience a differentially stronger house price boom (99-05) and a differentially stronger bust (07-09) in the high net-export-growth metropolitan areas, caused by growth of private-label mortgages (non-jumbo) 1999-2005. Second, after controlling the differential trend in APL-states, for all metropolitan areas, the growth of private-label mortgages (non-jumbo) 1999-2005 leads to a stronger house price boom (99-05) and a stronger bust (07-09) in the high net-export-growth areas.

Please note that in either boom or bust period, we have two endogenous variables here.

[70] This preemption is controversial afterward, especially since its spirit is somewhat against Section 1044 of Dodd-Frank (Sykes, 2019). However, in 2011, OCC reaffirmed the 2004 preemption while deleting some controversial languages from its preemption rules. OCC's formal document can be found here: `https://www.federalregister.gov/documents/2011/05/26/2011-12859/office-of-thrift-supervision-integration-dodd-frank-act-implementation`.

[71] According to Di Maggio and Kermani (2017), these 1eleven APL states (including DC) are California, Connecticut, District of Columbia, Georgia, Maryland, Michigan, Minnesota, New York, North Carolina, Texas, West Virginia.

$\triangle_{99,05} Ln(PLMNJ_c)$ is instrumented by $\triangle_{99,05}$givNetExp$_m$ and $\triangle_{99,05} Ln(PLMNJ_c) \times Dum_{APL}$ is instrumented by $\triangle_{99,05}$givNetExp$_m \times Dum_{APL}$. For each of the two separate first-stage regression F-tests, I report Sanderson-Windmeijer robust (clustered) F-statistics (Sanderson and Windmeijer, 2016). The SW F-statistics is 12.65 for PLMNJ growth and is 18.22 for the interaction between PLMNJ growth and dummy of APL-states in column (5), meaning each F-stage regression is significant and each instrument is strong for its endogenous variable. To evaluate the overall strength of two instruments, I report the p-value of robust (clustered) Kleibergen-Paap test statistic calculated by (Windmeijer, 2021). This p-value is 0.0078 in column (5), meaning the two separate instruments are jointly strong for the two endogenous variables.

A.4.2 Non-Recourse vs. Recourse States

In this subsection, I show that growth in private-label mortgages lead to stronger housing bust in non-recourse states. In addition, my main results are significant within the non-recourse states and within the recourse states.

Ghent and Kudlyak (2011) document that, in 39 states of the U.S., mortgages are recourse loans.[72] In these states, lenders could go after the borrower's other assets to recover the mortgage loss not covered by the proceedings from a foreclosure sale through obtaining a deficiency judgment. They find that in recourse states, borrowers are less sensitive to negative equity, and defaults are more likely to proceed via a lender-friendly procedure. I worry that non-recourse mortgages may discourage lenders from expanding mortgage credit but induce households to be more willing to default given falling house price.

Regression specification is the same as Eq (32) except that dummy variable is for non-recourse states. Table (A.4) reports the above 2SLS results and shows two key conclusion. First, compared to recourse states, non-recourse states experience a stronger house price bust (07-09), caused by growth of private-label mortgages (non-jumbo) 1999-2005. This evidence is consistent with the notion that non-recourse mortgages induce households to be more willing to default given falling house price. Second, after controlling the differential trend in the non-recourse states, for all metropolitan areas, the growth of private-label mortgages (non-jumbo) 1999-2005 leads to a stronger house price boom (99-05) and a stronger bust (07-09) in the high net-export-growth metropolitan areas. The SW F-statistics and p-value of robust (clustered) Kleibergen-Paap test statistics all show that instruments are both separately and jointly strong for the two endogenous variables.

[72] 39 recourse states can be found in Table 1 in Ghent and Kudlyak (2011). 11 non-recourse states are Alaska, Arizona, California, Iowa, Minnesota, Montana, North Carolina, North Dakota, Oregon, Washington, and Wisconsin.

A.4.3 Non-Judicial vs Judicial States

In this subsection, I show that growth in private-label mortgages lead to stronger housing bust in non-judicial states. In addition, my main results are significant within the non-judicial states and within the judicial states.

Mian, Sufi, and Trebbi (2015) shows that, in 20 states of the U.S., foreclosures of a delinquent property need judicial judgement.[73] In these states, in order to sell a delinquent property through foreclosure, lenders must file a notice with a judge providing evidence of the delinquency of the mortgage and get court approval. In contrast, in non-judicial states, the foreclosure process is much easier and does not need court approval. For more details, see Mian, Sufi, and Trebbi (2015). Mian, Sufi, and Trebbi (2015) find that lenders in non-judicial states are twice as likely to foreclose on delinquent mortgages. I worry that non-judicial states may experience stronger housing bust, caused by growth in private-label mortgages in the boom period (99-05).

Regression specification is the same as Eq (32) except that the dummy variable is for non-judicial states. Table (A.5) reports the above 2SLS results and shows two key conclusions. First, compared to judicial states, non-judicial states experience a stronger house price bust (07-09). This evidence is consistent with the notion that the absence of judicial procedure eases the process of foreclosure (Mian, Sufi, and Trebbi, 2015). Second, after controlling the differential trend in non-recourse states, for all metropolitan areas, the growth of private-label mortgages (non-jumbo) 1999-2005 leads to a stronger house price boom (99-05) and a stronger bust (07-09) in the high net-export-growth areas. The SW F-statistics and p-value of robust (clustered) Kleibergen-Paap test statistics all show that instruments are both separately and jointly strong for the two endogenous variables.

A.4.4 Sand vs Other States

In this subsection, I show that growth in private-label mortgages leads to a stronger housing price boom and bust cycle in sand states. In addition, my main results are significant within both the sand states and other states.

Many studies highlight that sand states (Arizona, California, Florida, and Nevada) experienced phenomenal housing cycles in comparison to the rest of the United States (Choi, Hong, Kubik, and Thompson, 2016). I expect that, in my setting, sand states may experience stronger housing bust, caused by growth in private-label mortgages in the boom period (99-05).

For regression specification, I add an interaction of dummy variable for sand states and pe-

[73] 20 judicial states can be found in Figure 2 in Mian, Sufi, and Trebbi (2015) and `https://www.realtytrac.com/real-estate-guides/foreclosure-laws/`. The twenty judicial states are Connecticut, Delaware, Florida, Illinois, Indiana, Kansas, Kentucky, Louisiana, Maine, Maryland, Massachusetts, Nebraska, New Jersey, New Mexico, New York, North Dakota, Ohio, Pennsylvania, South Carolina, and Vermont.

riod dummy to Eq (10). I do not use an interaction of three terms because the limited number of metropolitan areas and counties present a weak IV concern. These four sand states have 48 metropolitan areas and 73 counties. If I were to use the interaction of three terms $PLMNJGrowth(99-05) \times DumPeriod \times DumSand$, The SW F-statistics are 7.005 and 7.924 for the first-stage regression of PLMNJ Growth and $PLMNJGrowth \times DumSand$. Since here, my major focus is the differential housing boom and bust of sand states, a dummy variable can serve this purpose. In order to study the different impact from the growth of private-label mortgages, more granular data is necessary.

Table (A.6) reports the above 2SLS results and shows two key conclusions. First, compared to other states, sand states experience a stronger house price boom (99-05) and a stronger bust (07-09), shown in the coefficients of interaction terms of sand dummy and period dummy. Second, after controlling for the differential trend in the sand states, I only use the within-sand-states and within-other-states differences across metropolitan areas. The 2SLS regression shows that the growth of private-label mortgages (non-jumbo) 1999-2005 leads to a stronger house price boom (99-05) and a stronger bust (07-09) in the high net-export-growth areas. Since the cross-group differences between sand states and other states are removed by the interaction terms of sand dummy and period dummy, my 2SLS shall be interpreted as evidence strongly supporting my main conclusion: induced by net export growth, private-label mortgage causes the housing price boom (99-05) and a stronger bust (07-09) across all metropolitan areas. Meanwhile, the reduced cross-metro variation due to sand-state dummy reduces the F-statistics: the kleibergen-Paap (2006) robust (clustered) statistics is 7.362 and Montiel Olea-Pflueger (2013) efficient statistics is 7.302.

A.4.5 State Capital Gain Tax

In this section, I show that my major conclusion is robust to the inclusion of state capital gain tax as a control variable. Gao, Sockin, and Xiong (2020) find that speculation measured by non-owner-occupied purchase mortgages is discouraged by the state capital gain tax and such speculation contributes to the housing boom and bust. Based on their findings, I expect state capital gain tax to discourage housing boom and bust.

For regression specification, I add interaction of state capital gain tax rate and period dummy to Eq (10).[74] Table (A.7) reports the above 2SLS results and shows two key conclusions. First, the state capital gain tax rate contributes positively to the housing boom (99-05) and bust (07-09). This result contradicts with the discouraging effect of capital gain tax on housing speculation (Gao, Sockin, and Xiong, 2020). my result may reflect owner-occupied housing exclusion by state capital gain tax, which encourages owner-occupied home purchase. According to Tax Foundation[75],

[74] I would like to thank Zhenyu Gao, Michael Sockin, and Wei Xiong for sharing the state capital gain tax rate as of 2005.

[75] The Tax Foundation Report: `https://taxfoundation.org/research/all/federal/capital-gains-taxes/`.

homeowners may exclude up to $250,000 ($500,000 per couple) of capital gain if the homeowners had lived in the home for at least two of the previous five years. The exemption may be taken only once every two years and applies at both the federal and state level. Thus my results may sum the above two opposing effects of state capital gain on home purchase.

After controlling for state capital gain, my major conclusion holds: the growth of private-label mortgages (non-jumbo) 1999-2005 leads to the house price boom (99-05) and bust (07-09) across all metropolitan areas.

A.5 Mechanisms: Net Export Growth and Local Economy

In this section, I verify the assumption that net export growth causes a differential higher increase in local employment growth, household income growth, and population growth in the high net-export-growth metropolitan areas. This assumption is based on the economic base theory and links the causal relation between net export growth to credit expansion in private-label mortgages (PLM). The growth in employment, household income, and population are three channels through which net export growth induces credit expansion in PLM from 1999 to 2005. The analysis in this subsection includes all counties in the metropolitan areas (2003 CBSA code), since we do not rely on the HMDA mortgage data.

We do not claim these three channels are my contribution since there is ample evidence in the literature. Researchers show evidence of local employment changes due to imports (Autor, Dorn, and Hanson, 2013b; Pierce and Schott, 2016) and exports (Feenstra, Lipsey, Deng, Ma, and Mo, 2005). Studies also document the local household income changes due to imports (Barrot, Loualiche, Plosser, and Sauvagnat, 2022) and exports Minot and Goletti (1998). For changes in population growth (including migration), Greenland, Lopresti, and McHenry (2019) provides causal evidence on imports, and Li and Pan (2023) provides causal evidence on exports.

We conduct regression with the following specifications.

$$\triangle_{91,07} Ln(LocalEconConditions_c) = \beta * \triangle_{91,07} \text{NetExp}_m + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c \tag{33}$$

$\triangle_{91,07} Ln(LocalEconConditions_c)$ represents the growth of local economic condition at the county c, which can be employment, household income, and population. I use the gravity-model-based instrumental variable ($\triangle_{91,07}\text{givNetExp}_m$) as IV for $\triangle_{91,07}\text{NetExp}_m$. Since local sector employment composition and left-hand-side variables are possibly jointly determined, we only include basic controls, housing controls, and demographic controls. Weight and standard error are the same as equation (9).

A.5.1 Net Export and Employment

In this test, I use the total (excluding construction) employment share of the working-age population (age 15-64) because I want to avoid the impact of mortgage credit on construction employment between 1999 and 2007.

Table (A.8) reports OLS, reduced-form, second stage, and first stage of the results of total employment growth. First, panel A shows the positive and significant impact of net export growth (91-07, annualized) on total (excluding construction) employment share growth (91-07, annualized), without any control. Panel B reports the reduced-form estimates and shows that gravity-model-based net export growth increases household income growth. This significant impact is quite stable to the inclusion of various controls. First-stage estimates in panel D show that the strong positive correlation between the GIV and net export growth is quite stable across various specifications. The first-stage Kleibergen-Paap F-statistic (Kleibergen and Paap, 2006), which is robust to clustered standard errors, is 58.04 in column (4) with all control variables. And the Montiel Olea-Pflueger Efficient F-statistics (Olea and Pflueger, 2013) is 56.38 in column (4). As emphasized by Jiang (2017), it is important to compare the 2SLS estimates with OLS estimates given that 255 papers from the "Big Three" finance journals have the former one as nine times the latter one on average. The key concerns are the bias from a weak IV and publication bias in searching for "blown-up" IV estimates. Besides my two F-statistics, I also confirm that 2SLS estimates are all less than twice OLS estimates throughout various specifications, where my baseline results in column (4) indicate a ratio of $0.411/0.224 = 1.83$, much lower than nine. Thus, it is very unlikely that my estimates are biased by a weak instrument. We also show an added-variable plot of first-stage regression in figure (A.3) with full controls. We can see that the gravity model-based IV strongly predict net export growth with small standard error and high t-statistics (7.62).

Panel C reports 2SLS estimates. the 2SLS estimates are quite stable as I introduce the baseline controls in columns (2)-(4). According to column (4), one standard cross-sectional difference in annualized net export growth results in $0.411 \times 0.164\% = 0.067$ difference in employment share increase annually, translating into 1.076% difference from 1991 to 2007 at the metro level. Such an effect is statistically significant at the one percent level.

A.5.2 Net Export and Household Income

Table (A.9) reports OLS, reduced-form, second stage, and first stage of the results of household income growth. First, panel A shows the positive and significant impact of net export growth (91-07, annualized) on household income growth (91-07, annualized), with significance level at 1%. Similar pattern shows up in reduced-form estimates and 2SLS estimates. First-stage estimates in panel D show a strong positive correlation between the GIV and net export growth. Two first-stage

statistics are very high (with KP F-stat being 57.96 and MOP F-stat being 56.29 in column (4)). Additionally, the 2SLS estimates are very close to OLS estimates (column (4) indicates a ratio of $0.925/0.858 = 1.08$). Thus, my estimates are free from weak instrument concern.

As for economic meaning, 2SLS estimate in Column (4) in panel C shows one standard cross-sectional difference in annualized net export growth results in $0.925 \times 0.165\% = 0.153\%$ difference in household income growth, translating into 2.45% difference from 1991 to 2007 at the metro level.

A.5.3 Net Export and Population

Table (A.10) reports OLS, reduced-form, second stage, and first stage of the results of working-age population (age 15-64) growth. First, panel A shows the positive and significant impact of net export growth (91-07, annualized) on working-age population growth (91-07, annualized), all with significance level at 1%. Such a pattern shows up in reduced-form estimates and 2SLS estimates. First-stage estimates in panel D show a strong positive correlation between the net export growth and the gravity model-based instrumental variable. Two first-stage statistics are very high (with KP F-stat being 57.74 and MOP F-stat being 56.15 in column (4)). In addition, the 2SLS estimates are very close to OLS estimates, with column (4) indicating a ratio of $0.883/0.977 = 0.90$. Thus, my estimates are very unlikely biased by a weak instrument.

In terms of economic meaning, Column (4) in panel C shows one standard cross-sectional difference in annualized net export growth results in $0.883 \times 0.165\% = 0.146\%$ difference in working-age population growth, translating into 2.34% difference from 1991 to 2007 at the metro level.

A.6 Omitted Derivations in the Model

A.6.1 Boundary Conditions for income-to-house ratios

PLM in HNEG area

The assumption that all households can pay the full mortgage amount in up-state means that even a household with the lowest income can pay the full mortgage amount: $\underline{I_{H,0}}(1+g_u+\Delta g-c) \geq H_{H,0}(1+FCR_P^*+R_{P,H}+TF)$. This inequality implies

$$\frac{\underline{I_{H,0}}}{H_{H,0}} \geq \frac{1+FCR_P^*+R_{P,H}+TF}{1+g_u+\Delta g-c} := \underline{i_{P,H}}$$

The assumption that all households default in down-state means that even a household with the highest income could default: $\overline{I_{H,0}}(1+g_d+\Delta g-c) < H_{H,0}(1+FCR_P^*+R_{P,H}+TF)$. This inequality implies

$$\frac{\overline{I_{H,0}}}{H_{H,0}} < \frac{1+FCR_P^*+R_{P,H}+TF}{1+g_d+\Delta g-c} := \overline{i_{P,H}}$$

Therefore, I only assume that household have income-to-house ratio in time 0: $i_0 \in [\underline{i_{P,H}}, \overline{i_{P,H}})$.

This assumption matches the practice. For the households that would default with probability 1 ($i_0 < \underline{i_{P,H}}$ in my setting), lenders would not grant mortgages at all. For the households that would not default at all ($i_0 > \overline{i_{P,H}}$ in my setting), they could always get mortgages by paying the lowest rate (risk-free rate r_f), thus having no variation in mortgage rates. In either case, there is no economic trade-off in the mortgage market that is key to my model demonstration. Again, I purposefully ignore these two extreme groups of households to make my model simple and focus my attention on the key economic mechanism of interest.

PLM in LNEG area

The assumption that all households can pay the full mortgage amount in up-state means that even a household with the lowest income can pay the full mortgage amount: $\underline{I_{L,0}}(1+g_u-\Delta g-c) \geq H_{L,0}(1+FCR_P^*+R_{P,L}+TF)$. This inequality implies

$$\frac{\underline{I_{L,0}}}{H_{L,0}} \geq \frac{1+FCR_P^*+R_{P,L}+TF}{1+g_u-\Delta g-c} := \underline{i_{P,L}}$$

The assumption that all households default in down-state means that even a household with the highest income could default: $\overline{I_{L,0}}(1+g_d-\Delta g-c) < H_{L,0}(1+FCR_P^*+R_{P,L}+TF)$. This inequality implies

$$\frac{\overline{I_{L,0}}}{H_{L,0}} < \frac{1+FCR_P^*+R_{P,L}+TF}{1+g_d-\Delta g-c} := \overline{i_{P,L}}$$

Therefore, I only assume that household have income-to-house ratio in time 0: $i_0 \in [\underline{i_{P,L}}, \overline{i_{P,L}})$. This assumption matches the practice. For the households that would default with probability 1 ($i_0 < \underline{i_{P,L}}$ in my setting), lenders would not grant mortgages at all. For the households that would not default at all ($i_0 > \overline{i_{P,L}}$ in my setting), they could always get mortgages by paying the lowest rate (risk-free rate r_f), thus having no variation in mortgage rates. In either case, there is no economic trade-off in the mortgage market that is key to my model demonstration. Again, I purposefully ignore these two extreme groups of households to make my model simple and focus my attention on the key economic mechanism of interest.

A.6.2 When Up- and Down-states Have Unequal Probability

In the main part of the model, for easy exposition, I assume that both up- and down-states have the probability of $\frac{1}{2}$, as a way to model uncertainty at the household level. In a more general and realistic case, the up-state can have a probability $p > \frac{1}{2}$ (say around 95%). I will show in this appendix that my main prediction still holds that credit expansion in PLM is stronger in HNEG area than the one in LNEG area when PLM funding cost declines. This prediction means $i_{P,H}^{**} < i_{P,L}^{**}$ in math form.

First, let us derive $i_{P,H}^{**}$ for PLM in the HNEG area. The break-even condition for PLM in the

HNEG area shall be modified from Equation (17) as follows:

$$p * H_{H,0} \underbrace{(1 + FCR_P^* + R_{P,H} + TF)}_{MR_{P,H}^*} + (1-p)[\underbrace{RR * H_{H,0}(1 + \Delta g + \Delta e + \Delta p)}_{\text{Recovery Value}} + \underbrace{I_{H,0}(1 + g_d + \Delta g - c)}_{\text{Disposable Income}}] = H_{H,0}(1 + FCR_P^* + TF) \quad (34)$$

Then the mortgage rate $MR_{P,L}^*$ is set to make the equality hold:

$$MR_{P,H}^* = \frac{1}{p}[(1 + FCR_P^* + TF) - (1-p)RR(1 + \Delta g + \Delta e + \Delta p) - (1-p)(1 + g_d + \Delta g - c) * i_0] \quad (35)$$

Then I can derive the lower bound of the income-to-house ratio that can get PLM when the funding cost of PLM declines to FCR_P^{**}:

$$i_{P,H}^{**} = \frac{(1 + FCR_P^{**} + TF) - (1-p)RR(1 + \Delta g + \Delta e + \Delta p) - p\overline{MR_P}}{(1-p)(1 + g_d + \Delta g - c)} \quad (36)$$

Second, let us derive $i_{P,L}^{**}$ for PLM in the LNEG area. The break-even condition for PLM in the LNEG area shall be modified from Equation (19) as follows:

$$p * H_{L,0} \underbrace{(1 + FCR_P^* + R_{P,L} + TF)}_{MR_{P,L}^*} + (1-p)[\underbrace{RR * H_{L,0}(1 - \Delta g - \Delta e - \Delta p)}_{\text{Recovery Value}} + \underbrace{I_{L,0}(1 + g_d - \Delta g - c)}_{\text{Disposable Income}}] = H_{L,0}(1 + FCR_P^* + TF) \quad (37)$$

Then the mortgage rate $MR_{P,L}^*$ is set to make the equality hold:

$$MR_{P,L}^* = \frac{1}{p}[(1 + FCR_P^* + TF) - (1-p)RR(1 - \Delta g - \Delta e - \Delta p) - (1-p)(1 + g_d - \Delta g - c) * i_0] \quad (38)$$

Then I can derive the lower bound of the income-to-house ratio $i_{P,L}^{**}$ that can get PLM when the funding cost of PLM declines to FCR_P^{**}:

$$i_{P,L}^{**} = \frac{(1 + FCR_P^{**} + TF) - (1-p)RR(1 - \Delta g - \Delta e - \Delta p) - p\overline{MR_P}}{(1-p)(1 + g_d - \Delta g - c)} \quad (39)$$

By comparing both numerators and denominators, I can easily know $i_{P,H}^{**} < i_{P,L}^{**}$. Therefore, my major prediction still holds that credit expansion in PLM is stronger in HNEG area than one in LNEG area when PLM funding cost declines. Again, the necessary economic conditions for the above prediction are that there are additional positive growth rates in household income, employment, and net migration inflow in the high-net-export-growth areas.

A.7 Appendix Figures and Tables

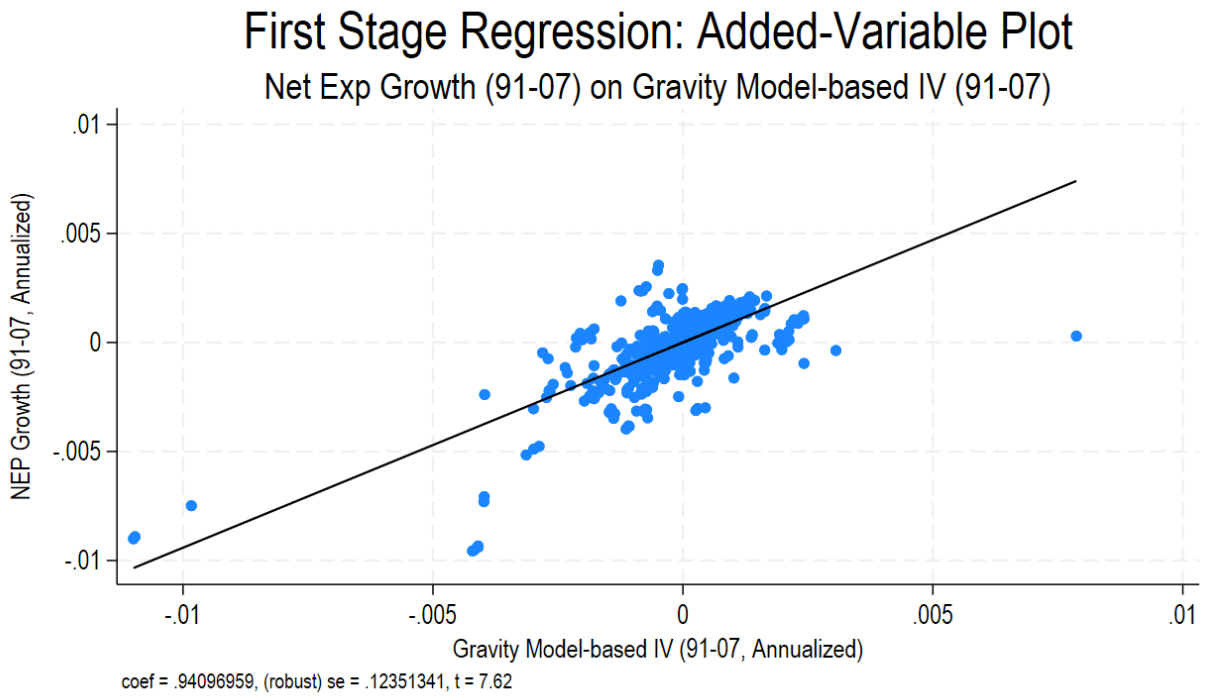


Figure A.3: **First Stage Regression: Net Export Growth (91-07) on Gravity Model-based Net Export Growth (91-07)**

This figure displays the added-variable plot of first stage regression: net export expansion growth (91-07) on gravity model-based instrumental variable growth (91-07) with full controls in column (4) in Table (A.8). Regression observations are weighted by logarithmic of number of total population in 1991 at county level. Standard errors are clustered at metropolitan areas (CBSA03) level. The coefficient is 0.941, clustered standard error is 0.124, and the t-statistic is 7.62.

Table A.3: **Robustness Test for Anti-Predatory Lending States vs. Non-Anti-Predatory Lending States.**
Stacked 2SLS Regressions of House Price Growth in Boom (99-05) and Bust (07-09) Periods on PLMNJ Growth (99-05)

This table reports OLS, reduced-form, first stage, and second stages of stacked 2SLS regression $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = \beta_{Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} + \beta_{APL,Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} \times Dum_{APL} + \beta_{APL,Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} \times Dum_{APL} + \gamma_{Boom} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{Bust} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c)$ is the stacked growth rate of the house price index at county c 99-05 and 07-09. The key variable of interest $\triangle_{99,05}Ln(PLMNJ_c$ is the growth rate of the dollar amount of private-label mortgages (non-jumbo) at county c 99-05. Dum_{APL} is the dummy variable for counties in states with anti-predatory lending laws. $Controls_c$ indicates control variables at county c in the period start year 1999. In either boom or bust period, we have two endogenous variables here: $\triangle_{99,05}Ln(PLMNJ_c)$ is instrumented by $\triangle_{99,05}\text{givNetExp}_m$ and $\triangle_{99,05}Ln(PLMNJ_c) \times Dum_{APL}$ is instrumented by $\triangle_{99,05}\text{givNetExp}_m \times Dum_{APL}$. For each of the first-stage F-tests of two endogenous variables, we report Sanderson-Windmeijer (2016) robust (clustered) statistics. To evaluate the overall strength of instruments, we report the p-value of robust (clustered) Kleibergen-Paap test statistics calculated by Windmeijer (2021). Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS Estimates	House Price Growth (07USD, 99-05 or 07-09, An)				
	(1)	(2)	(3)	(4)	(5)
PLMNJ Growth (99-05) x Dum99t05	0.432***	0.384***	0.368***	0.403***	0.384**
	(0.099)	(0.110)	(0.134)	(0.151)	(0.154)
PLMNJ Growth (99-05) x Dum07t09	-0.784***	-0.731***	-0.751***	-0.787***	-0.769***
	(0.178)	(0.189)	(0.254)	(0.242)	(0.241)
PLMNJ Growth (99-05) x Dum99t05 x DumAPL	0.166***	0.139***	0.103**	0.111**	0.097*
	(0.053)	(0.048)	(0.049)	(0.056)	(0.050)
PLMNJ Growth (99-05) x Dum07t09 x DumAPL	-0.288***	-0.256***	-0.246***	-0.218**	-0.227***
	(0.079)	(0.067)	(0.082)	(0.089)	(0.086)
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs	1572	1572	1390	1390	1390
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
SW F-Stat: PLMNJ (99 to 05) 1st-Stage	21.85	19.08	13.42	11.20	12.65
SW F-Stat: PLMNJxDumAPL (99 to 05) 1st-Stage	33.34	30.04	17.09	15.18	18.22
KP Robust (99 to 05) UnderID P-Value	0.0004	0.0004	0.0042	0.0074	0.0078
CoefEqual_Chi2	22.411	15.783	9.169	10.055	9.507
CoefEqual_PValue	0.000	0.000	0.002	0.002	0.002

Table A.4: **Robustness Test for Non-Recourse States vs. Recourse States.**
Stacked 2SLS Regressions of House Price Growth in Boom (99-05) and Bust (07-09) Periods on PLMNJ Growth (99-05)

This table reports stacked 2SLS regression $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = \beta_{Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} + \beta_{NRC,Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} \times Dum_{NRC} + \beta_{NRC,Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} \times Dum_{NRC} + \gamma_{Boom} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{Bust} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c)$ is the stacked growth rate of the house price index at county c 99-05 and 07-09 in either non-recourse state or recourse state. The key variable of interest $\triangle_{99,05}Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgages at the county c 99-05. Dum_{NRC} is the dummy variable for counties in non-recourse states. $Controls_c$ indicates control variables at county c in the start year 1999. In either boom or bust period, we have two endogenous variables here: $\triangle_{99,05}Ln(PLMNJ_c)$ is instrumented by $\triangle_{99,05}\text{givNetExp}_m$ and $\triangle_{99,05}Ln(PLMNJ_c) \times Dum_{APL}$ is instrumented by $\triangle_{99,05}\text{givNetExp}_m \times Dum_{APL}$. For each of the first-stage F-tests of two endogenous variables, we report Sanderson-Windmeijer (2016) robust (clustered) statistics. To evaluate the overall strength of instruments, we report the p-value of robust (clustered) Kleibergen-Paap test statistics calculated by Windmeijer (2021). Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS Estimates	House Price Growth (07USD, 99-05 or 07-09, An)				
	(1)	(2)	(3)	(4)	(5)
PLMNJ Growth (99-05) x Dum99t05	0.491***	0.433***	0.391***	0.431***	0.433**
	(0.110)	(0.119)	(0.135)	(0.155)	(0.176)
PLMNJ Growth (99-05) x Dum07t09	-0.880***	-0.816***	-0.808***	-0.836***	-0.837***
	(0.213)	(0.220)	(0.277)	(0.273)	(0.302)
PLMNJ Growth (99-05) x Dum99t05 x DumNRC	0.093	0.085	0.029	0.020	0.016
	(0.058)	(0.055)	(0.051)	(0.056)	(0.059)
PLMNJ Growth (99-05) x Dum07t09 x DumNRC	-0.230**	-0.236**	-0.229**	-0.193**	-0.200**
	(0.104)	(0.103)	(0.100)	(0.097)	(0.102)
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs	1572	1572	1390	1390	1390
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
SW F-Stat: PLMNJ (99 to 05) 1st-Stage	24.19	21.68	15.12	12.86	11.86
SW F-Stat: PLMNJxDumNRC (99 to 05) 1st-Stage	34.61	35.66	33.16	34.28	32.41
KP Robust (99 to 05) UnderID P-Value	0.0007	0.0007	0.0047	0.0084	0.0130
CoefEqual_Chi2	20.622	15.397	9.290	9.488	7.800
CoefEqual_PValue	0.000	0.000	0.002	0.002	0.005

Table A.5: **Robustness Test for Non-Judicial States vs. Judicial States.**
Stacked 2SLS Regressions of House Price Growth in Boom (99-05) and Bust (07-09) Periods on PLMNJ Growth (99-05)

This table reports stacked 2SLS regression $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = \beta_{Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} + \beta_{NJD,Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} \times Dum_{NJD} + \beta_{NJD,Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} \times Dum_{NJD} + \gamma_{Boom} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{Bust} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c)$ is the stacked growth rate of the house price index at county c 99-05 and 07-09 in either judicial state or non-judicial state. The key variable of interest $\triangle_{99,05}Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgages at the county c 99-05. Dum_{NJD} is the dummy variable for counties in states where foreclosure of a delinquent property needs judicial judgment. $Controls_c$ indicates control variables at county c in the start year 1999. In either boom or bust period, we have two endogenous variables here: $\triangle_{99,05}Ln(PLMNJ_c)$ is instrumented by $\triangle_{99,05}$givNetExp$_m$ and $\triangle_{99,05}Ln(PLMNJ_c) \times Dum_{APL}$ is instrumented by $\triangle_{99,05}$givNetExp$_m \times Dum_{APL}$. For each of the first-stage F-tests of two endogenous variables, we report Sanderson-Windmeijer (2016) robust (clustered) statistics. To evaluate the overall strength of instruments, we report the p-value of robust (clustered) Kleibergen-Paap test statistics calculated by Windmeijer (2021). Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS Estimates	House Price Growth (07USD, 99-05 or 07-09, An)				
	(1)	(2)	(3)	(4)	(5)
PLMNJ Growth (99-05) x Dum99t05	0.483***	0.409***	0.381***	0.425***	0.422***
	(0.102)	(0.109)	(0.131)	(0.149)	(0.163)
PLMNJ Growth (99-05) x Dum07t09	-0.842***	-0.745***	-0.739***	-0.774***	-0.793***
	(0.210)	(0.218)	(0.284)	(0.265)	(0.287)
PLMNJ Growth (99-05) x Dum99t05 x DumNJD	0.040	0.062*	0.024	0.016	0.023
	(0.043)	(0.032)	(0.027)	(0.029)	(0.028)
PLMNJ Growth (99-05) x Dum07t09 x DumNJD	-0.141**	-0.183***	-0.162***	-0.151***	-0.149***
	(0.058)	(0.051)	(0.054)	(0.054)	(0.055)
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs	1572	1572	1390	1390	1390
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
SW F-Stat: PLMNJ (99 to 05) 1st-Stage	21.28	19.13	13.45	11.36	11.22
SW F-Stat: PLMNJxDumNJD (99 to 05) 1st-Stage	45.87	51.44	37.91	36.91	37.06
KP Robust (99 to 05) UnderID P-Value	0.0011	0.0014	0.0091	0.0136	0.0150
CoefEqual_Chi2	20.932	14.309	8.155	9.362	8.360
CoefEqual_PValue	0.000	0.000	0.004	0.002	0.004

Table A.6: **Robustness Test for Sand States vs. Non-Sand States.**
Stacked 2SLS Regressions of House Price Growth in Boom (99-05) and Bust (07-09) Periods on PLMNJ Growth (99-05)

This table reports stacked 2SLS regression $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = \beta_{Boom} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{Bust} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} + \beta_{Sand,Boom} \times Dum_{99,05} \times Dum_{Sand} + \beta_{Sand,Bust} * \times Dum_{07,09} \times Dum_{Sand} + \gamma_{Boom} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{Bust} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_c$. The left-hand-side dependent variable $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c)$ is the stacked growth rate of the house price index at county c 99-05 and 07-09 in either sand state or non-sand state. The key variable of interest $\triangle_{99,05}Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgages at the county c 99-05. Dum_{Sand} is the dummy variable for counties in four sand states. $Controls_c$ indicates control variables at county c in the start year 1999. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05}Ln(PLMNJ_c)$. For the first-stage F-test of the non-stacked sample (99-05), we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS Estimates	House Price Growth (07USD, 99-05 or 07-09, An)				
	(1)	(2)	(3)	(4)	(5)
PLMNJ Growth (99-05) x Dum99t05	0.398***	0.353***	0.320**	0.366**	0.361**
	(0.120)	(0.124)	(0.133)	(0.162)	(0.176)
PLMNJ Growth (99-05) x Dum07t09	-0.613***	-0.564***	-0.533***	-0.555***	-0.628***
	(0.180)	(0.175)	(0.198)	(0.196)	(0.230)
DumSand x DumD99t05	0.040***	0.035**	0.029**	0.023	0.028**
	(0.014)	(0.014)	(0.013)	(0.015)	(0.014)
DumSand x DumD07t09	-0.112***	-0.109***	-0.111***	-0.102***	-0.097***
	(0.018)	(0.018)	(0.018)	(0.018)	(0.019)
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs	1572	1572	1390	1390	1390
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
KP F-Stat (99 to 05, non-stacked sample)	15.44	13.97	9.048	7.546	7.362
MOP F-Stat (99 to 05, non-stacked sample)	14.90	13.69	9.137	7.570	7.302
CoefEqual_Chi2	13.975	11.854	8.239	8.012	7.381
CoefEqual_PValue	0.000	0.001	0.004	0.005	0.007

Table A.7: **Robustness Test for State Capital Gain Tax Rates.**
Stacked 2SLS Regressions of House Price Growth in Boom (99-05) and Bust (07-09) Periods on PLMNJ Growth (99-05)

This table reports stacked 2SLS regression $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c) = \beta_{99,05} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{99,05} + \beta_{07,09} * \triangle_{99,05}Ln(PLMNJ_c) \times Dum_{07,09} + \beta_{Tax,Boom} \times StateCapGainTax_s \times Dum_{00,06} + \beta_{Tax,Bust} \times StateCapGainTax_s * \times Dum_{07,10} + \gamma_{99,05} * \boldsymbol{Controls_c} \times Dum_{99,05} + \gamma_{07,09} * \boldsymbol{Controls_c} \times Dum_{07,09} + \epsilon_{period,c}$. The left-hand-side dependent variable $\triangle_{99,05}\&\triangle_{07,09}Ln(HPI_c)$ is the stacked growth rate of the house price index at county c 99-05 and 07-09. The key variable of interest $\triangle_{99,05}Ln(PLMNJ_c)$ is the growth rate of the dollar amount of private-label mortgages at the county c 99-05. $StateCapGainTax_s$ is the 2005 state-level capital gain tax rate. $Controls_c$ indicates control variables at county c in the start year 1999. We add state-level capital gain tax rate interacted with period dummies. We use the gravity model-based instrumental variable ($\triangle_{99,05}$givNetExp$_m$) as IV for $\triangle_{99,05}Ln(PLMNJ_c)$. For the first-stage F-test of the non-stacked sample (99-05), we report kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of housing units in 1999. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

TSLS Estimates	House Price Growth (07USD, 99-05 or 07-09, An)				
	(1)	(2)	(3)	(4)	(5)
PLMNJ Growth (99-05) x Dum99t05	0.531***	0.479***	0.427***	0.478***	0.472***
	(0.112)	(0.118)	(0.146)	(0.170)	(0.178)
PLMNJ Growth (99-05) x Dum07t09	-0.937***	-0.887***	-0.864***	-0.920***	-0.963***
	(0.225)	(0.227)	(0.299)	(0.289)	(0.316)
State Capital Gain Tax x Dum99t05	0.242**	0.299***	0.229**	0.259**	0.239**
	(0.110)	(0.108)	(0.100)	(0.116)	(0.112)
State Capital Gain Tax x Dum07t09	-0.279	-0.408**	-0.375**	-0.430**	-0.468**
	(0.178)	(0.169)	(0.186)	(0.195)	(0.208)
DumPeriod	Y	Y	Y	Y	Y
Basic Controls x DumPeriod		Y	Y	Y	Y
Housing Controls x DumPeriod			Y	Y	Y
Demographic Controls x DumPeriod				Y	Y
Industry Controls x DumPeriod					Y
Obs	1572	1572	1390	1390	1390
Cluster SE	CBSA	CBSA	CBSA	CBSA	CBSA
Weight	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)	Ln(HU99)
KP F-Stat (99 to 05, non-stacked sample)	19.69	17.75	12.48	10.35	10.18
MOP F-Stat (99 to 05, non-stacked sample)	18.97	17.24	12.67	10.47	10.17
CoefEqual_Chi2	21.904	18.149	9.471	10.301	9.627
CoefEqual_PValue	0.000	0.000	0.002	0.001	0.002

Table A.8: **Four Regression of BEA Total (excluding construction) Employment Share Growth (91-07) on Net Export Growth (91-07)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{91,07} TotExConstEmpShr_c = \beta * \triangle_{91,07}\text{NetExp}_m + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{91,07} TotExConstEmpShr_c$ is the increase of BEA total (excluding construction) employment share per working-age population at county c 91-07 and the key variable of interest $\triangle_{91,07}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 91-07. $Controls_c$ indicates control variables at county c in 1991. We use the gravity model-based instrumental variable ($\triangle_{91,07}\text{givNetExp}_m$) as IV for $\triangle_{91,07}\text{NetExp}_m$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of the county-level population in 1991. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Total (ex const) Emp Share Growth (1991-2007, annualized)			
	(1)	(2)	(3)	(4)
Panel A. OLS estimates				
Net Export Growth (91-07, An)	0.392***	0.213**	0.238*	0.224*
	(0.097)	(0.105)	(0.124)	(0.118)
R2-adj	0.00975	0.0844	0.0867	0.124
Panel B. Reduced-form estimates				
GIV Net Export Growth (91-07, An)	0.531***	0.397***	0.407***	0.386***
	(0.137)	(0.115)	(0.136)	(0.142)
R2-adj	0.00880	0.0868	0.0888	0.126
Panel C . 2SLS estimates				
Net Export Growth (91-07, An)	0.493***	0.376***	0.433***	0.411***
	(0.102)	(0.101)	(0.127)	(0.132)
Dep Var (Panel D):	Net Export Growth (92-06, An)			
Panel D . First-stage estimates				
GIV Net Export Growth (91-07, An)	1.077***	1.057***	0.939***	0.941***
	(0.117)	(0.110)	(0.124)	(0.124)
KP F-Stat	85.18	92.75	57.44	58.04
MOP F-Stat	82.60	89.81	55.84	56.38
Controls (for all Panels)				
Basic Controls		Y	Y	Y
Housing Controls			Y	Y
Demographic Controls				Y
Obs	1041	1040	876	876
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(TotPop91)	Ln(TotPop91)	Ln(TotPop91)	Ln(TotPop91)

Table A.9: **Four Regressions of IRS Household Income Growth (91-07) on Net Export Growth (91-07)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{91,07}Ln(HHIncome_c) = \beta * \triangle_{91,07}\text{NetExp}_m + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{91,07}Ln(HHIncome_c)$ is the growth rate of the average household income (deflated to 2007 USD) at county c 91-07, and the key variable of interest $\triangle_{91,07}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 91-07. $Controls_c$ indicates control variables at county c in 1991. We use the gravity model-based instrumental variable ($\triangle_{91,07}\text{givNetExp}_m$) as IV for $\triangle_{91,07}\text{NetExp}_m$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of the number of households in 1991. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	IRS Household Income Growth (1991-2007, An, 07USD)			
	(1)	(2)	(3)	(4)
Panel A. OLS estimates				
Net Export Growth (91-07, An)	1.230***	1.049***	0.945***	0.858***
	(0.148)	(0.123)	(0.146)	(0.136)
R2-adj	0.0742	0.240	0.288	0.380
Panel B. Reduced-form estimates				
GIV Net Export Growth (91-07, An)	1.246***	1.081***	0.867***	0.875***
	(0.237)	(0.192)	(0.224)	(0.203)
R2-adj	0.0368	0.216	0.272	0.370
Panel C . 2SLS estimates				
Net Export Growth (91-07, An)	1.149***	1.016***	0.917***	0.925***
	(0.180)	(0.155)	(0.199)	(0.178)
Dep Var (Panel D):	Net Export Growth (91-07, An)			
Panel D . First-stage estimates				
GIV Net Export Growth (91-07, An)	1.085***	1.064***	0.945***	0.946***
	(0.117)	(0.111)	(0.125)	(0.124)
KP F-Stat	86.22	91.76	57.07	57.96
MOP F-Stat	83.78	88.74	55.40	56.29
Controls (for all Panels)				
Basic Controls		Y	Y	Y
Housing Controls			Y	Y
Demographic Controls				Y
Obs	1074	1074	894	894
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(NumHH91)	Ln(NumHH91)	Ln(NumHH91)	Ln(NumHH91)

Table A.10: **Four Regressions of Working-Age Population Growth (91-07) on Net Export Growth (91-07)**

This table reports OLS, reduced-form, first stage, and second stage results of 2SLS regression $\triangle_{91,07}Ln(WorkAgePop_c) = \beta * \triangle_{91,07}\text{NetExp}_m + \gamma * \boldsymbol{Controls_c} + \alpha + \epsilon_c$. The left-hand-side dependent variable $\triangle_{91,07}Ln(WorkAgePop_c)$ is the growth rate of the working-age population (age 15 - 64) at county c 91-07, and the key variable of interest $\triangle_{91,07}\text{NetExp}_m$ is the growth rate of net export at the metropolitan area (CBSA03 code) m 91-07. $Controls_c$ indicates control variables at county c in 1991. We use the gravity model-based instrumental variable ($\triangle_{91,07}\text{givNetExp}_m$) as IV for $\triangle_{91,07}\text{NetExp}_m$. For the first-stage F-test, we report Kleibergen-Paap (2006) robust (clustered) statistics and Montiel Olea-Pflueger (2013) efficient statistics. Each regression is weighted by the natural logarithm of the working-age population in 1991. Standard errors are clustered at the CBSA level. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Dep Var (Panel A, B, and C)	Working-Age Population Growth (1991-2007, annualized)			
	(1)	(2)	(3)	(4)
Panel A. OLS estimates				
Net Export Growth (91-07, An)	0.977***	1.156***	1.035***	0.977***
	(0.250)	(0.238)	(0.264)	(0.248)
R2-adj	0.0143	0.123	0.227	0.255
Panel B. Reduced-form estimates				
GIV Net Export Growth (91-07, An)	1.034***	1.182***	0.800**	0.835**
	(0.332)	(0.335)	(0.386)	(0.355)
R2-adj	0.00741	0.114	0.219	0.249
Panel C . 2SLS estimates				
Net Export Growth (91-07, An)	0.953***	1.110***	0.846**	0.883***
	(0.305)	(0.297)	(0.362)	(0.329)
Dep Var (Panel D):	Net Export Growth (91-07, An)			
Panel D . First-stage estimates				
GIV Net Export Growth (91-07, An)	1.085***	1.065***	0.945***	0.946***
	(0.117)	(0.111)	(0.125)	(0.124)
KP F-Stat	86.18	91.63	56.86	57.74
MOP F-Stat	83.65	88.63	55.27	56.15
Controls (for all Panels)				
Basic Controls		Y	Y	Y
Housing Controls			Y	Y
Demographic Controls				Y
Obs	1075	1074	894	894
Cluster SE	CBSA	CBSA	CBSA	CBSA
Weight	Ln(WorkAgePop91)	Ln(WorkAgePop91)	Ln(WorkAgePop91)	Ln(WorkAgePop91)